Global uncertainty and the dollar*

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Abstract

The US dollar tends to appreciate whenever global uncertainty goes up and vice versa. We assess whether a) this co-movement is caused by global uncertainty shocks and—to the extent that it is—whether b) the 'dollar channel' matters for their global repecussions. In a first step, we identify global uncertainty shocks in a time-series model and establish that they cause a synchronized slowdown of the world economy as well as an appreciation of the dollar. They also appreciate other safe-haven currencies. Second, we construct counterfactual scenarios in which the dollar is unresponsive to global uncertainty shocks and find their contractionary effect outside of the US much reduced—in contrast to what we find for other safe-haven currencies. This testifies to the special role of the dollar in the transmission of global uncertainty shocks.

Keywords: US dollar, safe-haven currencies, uncertainty shocks, trade channel, financial channel, Bayesian Proxy VAR, minimum relative entropy, counterfactual, monetary policy

JEL-Classification: F31, F42, F44

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1 Introduction

The Global Financial Crisis (GFC) and the COVID-19 pandemic share a common feature: economic uncertainty reached record levels in the context of both events. The black solid lines in Figure 1 depict the expected stock market volatility index (VIX) compiled by the Chicago Board of Options—a widely-used measure of uncertainty—around both events. But there is a second noteworthy observation: the US dollar exchange rate—depicted by the red dashed lines—also appreciated sharply in both instances. In fact, the strong co-movement of the VIX and the dollar is a systematic pattern in the data that is not confined to the GFC and the COVID-19 pandemic.[1] It is also a direct implication of dollar's role as a safe-haven or reserve currency (Maggiori 2017; He et al. 2019; Jiang et al. forthcoming), and the notion that the 'exorbitant privilege' of the US comes with an 'exorbitant duty', namely adverse valuation effects on its external balance sheet in times of crisis (Gourinchas & Rey 2007; Gourinchas et al. 2010).

In this paper, we first document that global uncertainty shocks cause a positive co-movement of the VIX and the dollar. In line with earlier work, we find that uncertainty shocks trigger a contraction of economic activity (Fernández-Villaverde et al. 2015; Baker et al. 2016; Basu & Bundick 2017). Importantly, we also show that the contractionary effects are highly synchronized across the US and the rest of the world, and that they induce a monetary policy accommodation across the globe. Moreover, we establish that global uncertainty shocks, while inducing a rise in the VIX, cause a sharp and persistent appreciation of the dollar exchange rate; other currencies widely seen as safe havens such as the Japanese yen and the Swiss franc appreciate as well. In contrast, other currencies—such as the euro or the British pound—depreciate.

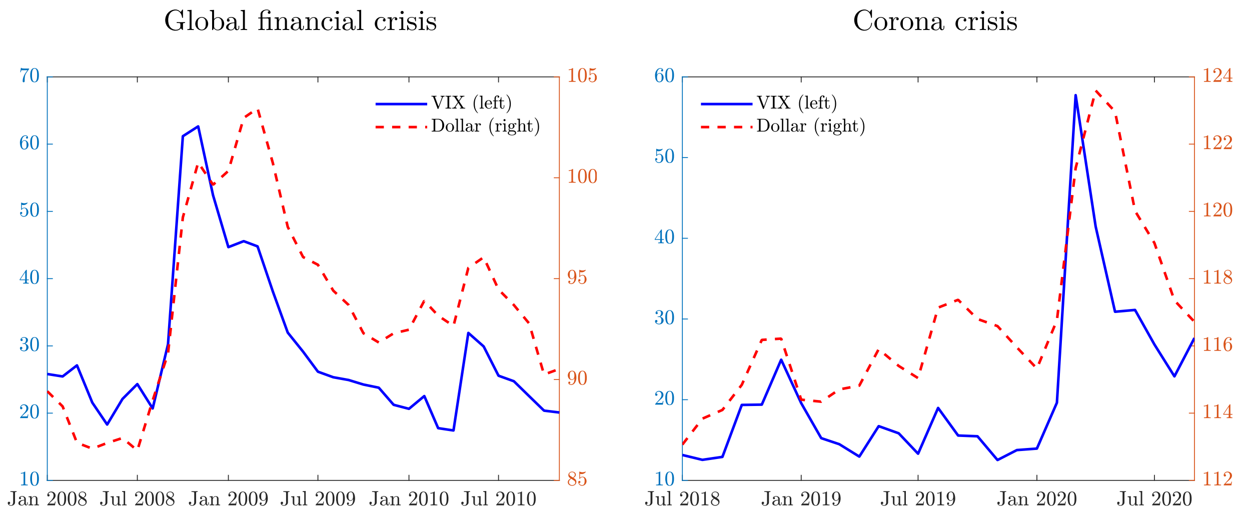
In a second step, we ask whether the appreciation of the dollar matters for the transmission of global uncertainty shocks to the rest of the world. In theory, the effect of dollar appreciation is ambiguous. On the one hand, dollar appreciation dampens the adverse impact of global uncertainty shocks in the rest of the world via a 'trade channel', as it induces expenditure switching of demand from the US towards the rest of the world (Obstfeld & Rogoff 1996).[2] On the other hand, dollar appreciation amplifies the adverse impact of global uncertainty shocks in the rest of the world via a 'financial channel', as it induces a contraction in cross-border bank credit (Bruno & Shin 2015).

Which of these two competing channels dominates and hence whether dollar appreciation dampens or amplifies the effects of global uncertainty shocks is an empirical question. In order to address this question we examine a counterfactual and simulate the effects of a global uncertainty shock that would materialise in the absence of a dollar appreciation. We find that the financial channel dominates the trade channel: In the counterfactual US net exports and—especially US dollar-denominated—cross-border bank credit to non-US borrowers contract less in response to the uncertainty shock; all else equal, the first implies a harsher slowdown of economic activity in the rest

[1] The t-value in a monthly regression of first differences of the VIX on logarithmic first differences of the dollar exchange rate over the period from 01/2003 to 04/2020 is 6.7 and 3.5 when excluding 10/2008 and 03/2020.

[2] While this mechanism was originally put forth in the Mundellian paradigm in which export prices are sticky in the currency of the producer, it is still present at least for US exports under the dominant-currency paradigm when trade prices are sticky in dollar (Gopinath et al. 2020).

Figure 1: The US dollar exchange rate and the VIX



Notes: VIX is index of expected stock market volatility compiled by Chicago Board of Options Exchange (source: Wall Street Journal/Haver); dollar exchange rate is price of dollar expressed in foreign currency (in effective terms) such that an increase represents an appreciation (source: Federal Reserve Board/Haver).

of the world, while the second implies a weaker slowdown. Overall, we find that economic activity in the rest of the world contracts considerably less in the counterfactual.

In more detail, we estimate a structural vector autoregressive (SVAR) model with time series for the VIX, industrial production in the US and the rest of the world, the US consumer price index and excess bond premium, the one-year Treasury Bill rate as an indicator of US monetary policy, rest-of-the-world policy rates, and the US dollar effective exchange rate. In addition, we also include one at a time time series for US trade and cross-border bank credit flows. We estimate the VAR model on 30 years of monthly observations, covering the period from 1990 to 2019.

To achieve identification we rely on proxy variables as instruments (Stock & Watson 2012; Mertens & Ravn 2013). We adopt the Bayesian proxy SVAR framework of Arias et al. (forthcoming), which allows us to refrain from imposing restrictions on the contemporaneous relationships between endogenous variables. Instead we use an external instrument ("proxy variable") to identify global uncertainty shocks: the intra-daily changes of the gold price in a narrow window around events commonly believed to represent exogenous changes of global uncertainty (Piffer & Podstawski 2018; Engel & Wu 2018; Ludvigson et al. forthcoming). We use a second external instrument to identify a US monetary policy shock, namely the high-frequency interest rate changes in a narrow window around Federal Open Market Committee announcements (Gertler & Karadi 2015; Caldara & Herbst 2019; Jarociński & Karadi 2020). The Bayesian proxy SVAR framework of Arias et al. (forthcoming) is suited particularly well for our purposes, as it allows for coherent estimation and inference when multiple structural shocks are jointly identified by means of external instruments.

Our first set of result concerns the effects of a global uncertainty shock. We that a one-standard-deviation shock appreciates the dollar by about 0.5 percent, industrial production drops by 0.4

percent both in the US and the rest of the world, and short-term policy rates decline by 6 to 8 basis points. US exports and imports contract by about 0.5 percent; consistent with dominant-currency paradigm (DCP; Gopinath et al. 2020) US exports contract already in the short term, while US imports respond only sluggishly. Cross-border bank credit to non-US borrowers contracts by about 1 percent. The Japanese Yen appreciates by about 1 percent and the Swiss Franc by 0.2 percent.

In order to evaluate the role of the exchange rate for the transmission of global uncertainty shocks—notably in the rest of the world—we examine counterfactuals in which the dollar is unresponsive. We pursue two alternative approaches to develop the counterfactuals. The first approach is agnostic regarding the structural factors that render the dollar exchange rate unresponsive to global uncertainty shocks; the second approach represents a policy experiment.

Consider the agnostic counterfactual first. It is based on the concept of 'minimum relative entropy' (MRE) developed in the context of forecasting (Robertson et al. 2005; Cogley et al. 2005; Giacomini & Ragusa 2014). Intuitively, the idea is to improve forecasts by imposing a restrictions derived from theory. Following Breitenlechner et al. (2020) we apply the MRE approach in computing impulse responses. We think of these as conditional forecasts and in the context of counterfactual analysis: we construct counterfactual responses to a global uncertainty shock which are characterized by a) the absence of dollar appreciation while b) being as similar to the baseline as possible otherwise (hence the name MRE). Our second set of results are based on these counterfactuals. We find that When the dollar is unresponsive to a global uncertainty shock, US consumer prices and policy rates decline less in response to a global uncertainty shock. The same holds for US net exports. A smaller contraction of US net exports, all else equal, amplifies the contractionary effect of a global uncertainty shock in the rest of the world. However, in the counterfactual the decline in cross-border bank credit to non-US borrowers is also less pronounced. And a smaller drop in cross-border bank credit, all else equal, dampens the contractionary effect of a global uncertainty shock in the rest of the world. Overall, when dollar appreciation is absent the contractionary effects of a global uncertainty shock on rest-of-the-world industrial production are almost halved compared to the baseline. Our results thus show that the financial channel dominates the trade channel. Moreover, we find a special role of the US dollar for the transmission of global uncertainty shocks: if we construct a counterfactuals in which the Japanese yen and the Swiss franc, in turn, are unresponsive the effect of the uncertainty shock is virtually unchanged.

The second counterfactual represents a policy experiment and is based on the materialisation of additional shocks over time (Kilian & Lewis 2011; Bachmann & Sims 2012; Wong 2015; Epstein et al. 2019). Specifically, we assume a sequence of US monetary policy shocks that offsets the effect of the global uncertainty shock on the dollar exchange rate. We refer to this as a 'structural shock counterfactual' (SSC; Antolin-Diaz et al. 2021). In other words, we explore how a global uncertainty shock would play out if US monetary policy—in contrast to the regularities in the data—intervened to stabilise the dollar exchange rate. We find, and this is our third result, that by adopting a more accommodative stance that stabilises the dollar exchange rate US monetary policy would mitigate substantially the contractionary effects of global uncertainty shocks, both in the US and in the rest of the world. In fact, the results for the structural shock counterfactual are very similar to those for counterfactual based on the MRE criterion. And while they are conceptually and methodologically quite distinct, both counterfactuals illustrate the special role of the US dollar for the transmission of global uncertainty shocks.

Our paper relates to the literature on the special role of the US and the dollar for the world economy and the global financial cycle (Rey 2016; Miranda-Agrippino & Rey 2020). Cerutti et al. (2017), Avdjiev, Bruno, et al. (2019) as well as Avdjiev, Du, et al. (2019) provide evidence for a special role of the dollar exchange rate in the transmission of the global financial cycle to local credit supply. Typically, this literature focuses on the implications of US and/or global shocks for monetary policy and financial conditions in the rest of the world, in particular in emerging market economies (Alfaro et al. 2019; Kalemli-Özcan 2019; Vicondoa 2019). However, there is also evidence that dollar appreciation tightens financing conditions in the US itself (Niepmann & Schmidt-Eisenlohr 2017; Meisenzahl et al. 2019).

There is also closely related work on the international spillovers of credit and uncertainty shocks (Carriere-Swallow & Cespedes 2013; Cesa-Bianchi et al. 2018; Epstein et al. 2019; Bhattarai et al. 2020) as well as US monetary policy shocks (Georgiadis 2016; Dedola et al. 2017; Dees & Galesi 2019; Iacoviello & Navarro 2019; Degasperi et al. 2020). In addition, there are more reduced-form approaches, which investigate the effects of 'dollar shocks'—which lump together uncertainty, monetary policy, demand and other shocks—on the global economy (Liu et al. 2017; Shousha 2019).

Lastly, there is important work on how exchange rates contribute to the cross-border transmission of shocks via the financial channel. Banerjee et al. (2016), Aoki et al. (2018) as well as Akinci & Queralto (2019) put forth theoretical models in which dollar appreciation amplifies the cross-border transmission of shocks due to currency mismatches on borrowers' balance sheets. Shim et al. (2020) show that firms in 10 emerging market economies whose non-financial sectors hold more debt in foreign currency decrease their leverage relatively more after home currency depreciations. Hofmann et al. (2016, 2019) document empirically that the financial channel of exchange rates is not confined to private-sector, but also impacts sovereign bond borrowing costs. Using firm-level data for 18 major global economies, Banerjee et al. (2020) find that exchange rate depreciation dampens corporate investment through firm leverage and foreign-currency debt. Bruno & Shin (2019) provide empirical evidence that dollar appreciation inhibits exports for firms which rely on dollar liquidity to finance working capital. Kearns & Patel (2016) also study whether the financial channel offsets the trade channel, but they do not consider identified shocks.

The rest of the paper is organised as follows. Section 2 discusses identification in a Bayesian proxy SVAR framework with multiple proxies. Section 3 presents the empirical specification of the Bayesian proxy SVAR we bring to the data. Section 4 presents our findings for impulse responses to global uncertainty and US monetary policy shocks as well as the MRE counterfactuals and the SSC. Finally, Section 5 concludes.

2 The Bayesian proxy structural VAR model

In this section we lay out the BPSVAR model of Arias et al. (forthcoming) that forms our empirical framework. Providing a general description allows us to highlight some appealing features of the BPSVAR framework relative to previous approaches for the identification of uncertainty shocks and the estimation of their macroeconomic effects.

Following the notation of Rubio-Ramirez et al. (2010), consider without loss of generality the structural VAR model with one lag and without deterministic terms

$$\boldsymbol{y}_t'\boldsymbol{A}_0 = \boldsymbol{y}_{t-1}'\boldsymbol{A}_1 + \boldsymbol{\epsilon}_t', \tag{1}$$

where $\boldsymbol{y}_t$ is an $n \times 1$ vector of endogenous variables and $\boldsymbol{\epsilon}_t$ an $n \times 1$ vector of structural shocks. The BPSVAR framework builds on the following assumptions in order to identify k structural shocks of interest: There exists a $k \times 1$ vector of proxy variables $\boldsymbol{m}_t$ that are (i) correlated with the k structural shocks of interest $\boldsymbol{\epsilon}_t^*$, and (ii) orthogonal to the remaining structural shocks $\boldsymbol{\epsilon}_t^o$. Formally, the identifying assumptions are

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{*\prime}] = \underset{(k\times k)}{\boldsymbol{V}}, \tag{2a}$$

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{o\prime}] = \underset{(n-k\times k)}{\boldsymbol{0}}, \tag{2b}$$

and are known as the relevance and the exogeneity condition, respectively.

Given Equation (1) as well as Equations (2a) and (2b), Arias et al. (forthcoming) augment the model in Equation (1) with k proxy equations. In particular, denote by $\tilde{\boldsymbol{y}}_t' \equiv (\boldsymbol{y}_t', \boldsymbol{m}_t')$ the vector of endogenous variables augmented with the $k \times 1$ vector of proxy variables, by $\tilde{\boldsymbol{A}}_\ell$ the corresponding coefficient matrices of dimension $\tilde{n} \times \tilde{n}$ with $\tilde{n} = n + k$, by $\tilde{\boldsymbol{\epsilon}} \equiv (\boldsymbol{\epsilon}_t', \boldsymbol{v}_t')' \sim N(\boldsymbol{0}, \boldsymbol{I}_{n+k})$, where $\boldsymbol{v}_t$ is a $k \times 1$ vector of measurement errors that affect the proxy variables (see below). The augmented model is then given by

$$\tilde{\boldsymbol{y}}_t'\tilde{\boldsymbol{A}}_0 = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 + \tilde{\boldsymbol{\epsilon}}_t'. \tag{3}$$

To ensure that augmenting the model with equations for the proxy variables does not affect the dynamics of the endogenous variables, restrictions are imposed on the matrices $\tilde{\boldsymbol{A}}_\ell$ such that

$$\tilde{\boldsymbol{A}}_\ell = \begin{pmatrix} \underset{(n\times n)}{\boldsymbol{A}_\ell} & \underset{(n\times k)}{\boldsymbol{\Gamma}_{\ell,1}} \\ \underset{(k\times n)}{\boldsymbol{0}} & \underset{(k\times k)}{\boldsymbol{\Gamma}_{\ell,2}} \end{pmatrix}, \qquad \ell = 0, 1. \tag{4}$$

The zero restrictions on the lower left-hand side block imply that the proxy variables do not enter the equations of the endogenous variables. The reduced form of the model is

$$\tilde{\boldsymbol{y}}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1\tilde{\boldsymbol{A}}_0^{-1} + \tilde{\boldsymbol{\epsilon}}_t{}'\tilde{\boldsymbol{A}}_0^{-1}. \tag{5}$$

Because the inverse of $\tilde{\boldsymbol{A}}_0$ is given by

$$\tilde{\boldsymbol{A}}_0^{-1} = \begin{pmatrix} \boldsymbol{A}_0^{-1} & -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ 0 & \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix}, \tag{6}$$

the last k equations of Equation (5) read as

$$\boldsymbol{m}_t' = \tilde{\boldsymbol{y}}_{t-1}'\tilde{\boldsymbol{A}}_1 \begin{pmatrix} -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} \\ \boldsymbol{\Gamma}_{0,2}^{-1} \end{pmatrix} - \boldsymbol{\epsilon}_t'\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} + \boldsymbol{v}_t'\boldsymbol{\Gamma}_{0,2}^{-1}, \tag{7}$$

which shows that the proxy variables may be serially correlated, affected by past values of the endogenous variables and by measurement error.

Ordering the structural shocks as $\boldsymbol{\epsilon}_t = (\boldsymbol{\epsilon}_t^{o\prime}, \boldsymbol{\epsilon}_t^{*\prime})'$ we have

$$E\left[\boldsymbol{m}_t\boldsymbol{\epsilon}_t'\right] = -\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1} = \begin{pmatrix} \underset{(k\times(n-k))}{\boldsymbol{0}} & \underset{(k\times k)}{\boldsymbol{V}} \end{pmatrix}, \tag{8}$$

where the first equality is obtained using Equation (7) and because the structural shocks $\boldsymbol{\epsilon}_t$ are by assumption orthogonal to $\boldsymbol{y}_{t-1}$ and $\boldsymbol{v}_t$, and the second equality is due to the exogeneity and relevance conditions in Equations (2a) and (2b). Equation (8) shows that the identifying assumptions imply restrictions on the last k columns of the contemporaneous structural impact coefficients in $\tilde{\boldsymbol{A}}_0^{-1}$. In particular, if the exogeneity condition in Equation (2b) holds, the first $n-k$ columns of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ in Equation (6) are zero. From Equation (5) it can be seen that this implies that the first $n-k$ structural shocks do not impact contemporaneously the proxy variables. In turn, if the relevance condition in Equation (2a) holds, the last k columns of the upper right-hand side sub-matrix $\boldsymbol{A}_0^{-1}\boldsymbol{\Gamma}_{0,1}\boldsymbol{\Gamma}_{0,2}^{-1}$ of $\tilde{\boldsymbol{A}}_0^{-1}$ are different from zero. From Equation (5) it can be seen that this implies that the last k structural shocks impact the proxy variables contemporaneously. In the algorithm of Arias et al. (forthcoming) the estimates of $\boldsymbol{A}_0$ and $\boldsymbol{\Gamma}_{0,\ell}$ are obtained such that the restrictions on $\tilde{\boldsymbol{A}}_0^{-1}$ implied by Equations (2a) and (2b) are satisfied, and hence the estimation identifies the structural shocks of interest in $\boldsymbol{\epsilon}_t^*$.

If the number of structural shocks identified by the proxy variables is larger than one, the relevance and exogeneity conditions are not sufficient for point identification (Giacomini et al. forthcoming), as rotations of the structural shocks $\boldsymbol{Q}\boldsymbol{\epsilon}_t^*$ also satisfy the exogeneity and relevance conditions in Equations (2a) and (2b). In this case, additional restrictions are needed in order to point-identify the structural shocks in $\boldsymbol{\epsilon}_t^*$. In the traditional proxy SVAR model these additional restrictions are imposed on the contemporaneous relationships between the endogenous variables in $\boldsymbol{A}_0^{-1}$ (Mertens & Ravn 2013; Lakdawala 2019). However, Arias et al. (forthcoming) show that relaxing this type of additional identifying assumptions can drastically change the results. An advantage of the BPSVAR framework is that the additional identifying assumptions can instead be imposed on the relevance condition in Equation (2a) reflected in the matrix $\boldsymbol{V}$. For example, one can impose the restriction that a particular structural shock does not affect a particular proxy variable.[3] Restrictions on the

[3] Jarociński & Karadi (2020) identify a monetary policy shock and a central bank information shock with two proxy

relationship between structural shocks and proxy variables in $\boldsymbol{V}$ are arguably less controversial and hence weaker than exogeneity assumptions on the relationship between the endogenous variables.

Another appealing feature of the BPSVAR model is that it allows to incorporate a prior belief about the strength of the proxy variables as instruments based on the notion that "researchers construct proxies to be relevant" (Caldara & Herbst 2019, p. 165). A convenient metric is the 'reliability matrix' $\boldsymbol{R}$ derived in Mertens & Ravn (2013) given by

$$\boldsymbol{R} = \left(\boldsymbol{\Gamma}_{0,2}^{-1\prime}\boldsymbol{\Gamma}_{0,2} + \boldsymbol{V}\boldsymbol{V}'\right)^{-1}\boldsymbol{V}\boldsymbol{V}'. \tag{9}$$

Intuitively, $\boldsymbol{R}$ indicates the share of variance of the proxy variables that is accounted for by the structural shocks $\boldsymbol{\epsilon}_t^*$ in their total variance (see Equation (7)). Specifically, the minimum eigenvalues of $\boldsymbol{R}$ can be interpreted as the share of the variance of (any linear combination of) the proxy variables explained by the structural shocks $\boldsymbol{\epsilon}_t^*$ (Gleser 1992).

Especially relative to frequentist proxy SVAR frameworks another appealing feature of the BPSVAR framework is that inference is straightforward, both in case of set and point identification with and without sign, magnitude and zero restrictions. This is an important issue, as—to the best of our knowledge—there is no consensus yet on how to conduct inference in frequentist proxy SVAR models, even in a setting with only a single proxy variable (Jentsch & Lunsford 2016, 2019). Inference is even more challenging when the proxy variable is considered to be a weak instrument (Olea et al. forthcoming), or when the identifying assumptions achieve only set rather than point identification (Moon & Schorfheide 2012).

Finally, yet another appealing feature of the BPSVAR model is that it allows to identify additional structural shocks using zero, sign and magnitude restrictions. To do so, as in a traditional SVAR framework additional restrictions may be imposed on the contemporaneous structural impact matrix $\boldsymbol{A}_0^{-1}$. The BPSVAR framework allows rigorous inference for specifications that mix identification with zero, sign, and magnitude restrictions as well as proxy variables.

To sum up, several considerations render the BPSVAR model of Arias et al. (forthcoming) particularly appealing in order to estimate the effects of global uncertainty shocks: (i) the possibility to avoid recursiveness assumptions for the identification of uncertainty shocks, (ii) our requirement to jointly identify global uncertainty and US monetary policy shocks, and (iii) the possibility to carry out coherent inference when identification is achieved by (multiple) proxy variables.[4]

variables. However, they do so by including the proxy variables as internal instruments in the VAR model and then impose two sets of sign restrictions on the responses of endogenous variables, thus only achieving set identification.

[4]An issue that we do not address in this paper is the possibility of a non-linear relationship between uncertainty shocks one the one hand and the VXO and the dollar on the other hand. We leave this for future research.

3 VAR specification and identification assumptions

3.1 VAR specification

Our point of departure is the US VAR model of Gertler & Karadi (2015) which includes in $\boldsymbol{y}_t$ the logarithms of US industrial production and consumer prices, the excess bond premium of Gilchrist & Zakrajsek (2012), and the one-year Treasury Bill rate as monetary policy indicator. We augment $\boldsymbol{y}_t$ with the VXO, the logarithm of an index of non-US, rest-of-the-world (RoW) industrial production, a weighted average of advanced economies' (AEs') policy rates, and the logarithm of the US dollar nominal effective exchange rate (NEER).[5] We use monthly data for the time period from February 1990 to June 2019. Data descriptions are provided in Table A.1.

3.2 Proxy variables

Following the literature on high-frequency identification (Kuttner 2001; Gürkaynak et al. 2005) we consider the intra-daily gold price changes of Piffer & Podstawski (2018) around narrow time windows on narratively selected days as proxy variable for the global uncertainty shock. In particular, Piffer & Podstawski (2018) first extend the list of dates selected by Bloom (2009) on which the VXO increased arguably due to exogenous uncertainty shocks. Second, they calculate the change in the price of gold between the last auction before and the first auction after the news about the event representing the uncertainty shock became available to markets. The analysis of Piffer & Podstawski (2018) covers the time period until 2015; we use the update of Bobasu et al. (2020) that spans until 2019.

As in Gertler & Karadi (2015) as well as Caldara & Herbst (2019) we consider intra-daily 3-month Federal Funds Futures rate changes around narrow time windows on FOMC meeting days of Gürkaynak et al. (2005) as proxy variable for the US monetary policy shock. We in addition cleanse these interest rate surprises from central bank information effects using the poor-man's approach of Jarociński & Karadi (2020): When the Federal Funds Futures rate surprise around an FOMC announcement has the same sign as the equity price surprise, this is classified as central bank information effect; when the Federal Funds Futures rate and the equity price surprises have the opposite sign, this is classified as 'pure' monetary policy surprise.[6]

Figure A.1 plots the monthly time series of the gold price and Federal Funds Futures rate surprises as we use them as proxy variables in the estimation of our BPSVAR model. The biggest (positive) spikes in the gold price surprise are recorded for the launch of Operation Desert Storm in the early 1990s, the 9/11 attacks in 2001, when American International Group (AIG) requested emergency lending at the height of the Global Financial Crisis in 2008, and the release of the results

[5]We use AE instead of RoW policy rates as the latter exhibit dramatic spikes stemming from periods of hyperinflation in some EMEs.

[6]We aggregate the daily changes in the proxy variables to monthly frequency as in Gertler & Karadi (2015). In particular, we first create a cumulative daily surprise series, then, second, take monthly averages of these series, and, third, obtain monthly average surprises as the first difference of this series. Note that while this may induce serial correlation in the interest rate surprises, this is allowed for in the BPSVAR framework (see Equation (7)).

of the Brexit referendum in 2016.

3.3 Identifying assumptions

Define $\boldsymbol{\epsilon}_t^* \equiv (\epsilon_t^u, \epsilon_t^{mp})'$, where ϵ_t^u denotes the global uncertainty shock and ϵ_t^{mp} the US monetary policy shock. Furthermore, define $\boldsymbol{m}_t = (p_t^{\epsilon,u}, p_t^{\epsilon,mp})'$ as the vector containing the proxy variables for the global uncertainty and the US monetary policy shock, that is the gold price and Federal Funds Futures surprises.

Our identifying assumptions are given by

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{*\prime}] = \begin{pmatrix} E[p_t^{\epsilon,u}\epsilon_t^u] & E[p_t^{\epsilon,u}\epsilon_t^{mp}] \\ E[p_t^{\epsilon,mp}\epsilon_t^u] & E[p_t^{\epsilon,mp}\epsilon_t^{mp}] \end{pmatrix} = \boldsymbol{V}, \tag{10a}$$

$$E[\boldsymbol{m}_t\boldsymbol{\epsilon}_t^{o\prime}] = \begin{pmatrix} E[p_t^{\epsilon,u}\boldsymbol{\epsilon}_t^o] & E[p_t^{\epsilon,mp}\boldsymbol{\epsilon}_t^o] \end{pmatrix} = \mathbf{0}. \tag{10b}$$

First, in the relevance condition in Equation (10a) we assume that global uncertainty shocks drive the gold price surprises on the narratively selected dates, $E[p_t^{\epsilon,u}\epsilon_t^u] \neq 0$. Intuitively, as gold is widely seen as a safe haven asset, when uncertainty rises increases in precautionary savings push up the price of gold (Baur & McDermott 2010, 2016). Piffer & Podstawski (2018) provide evidence that gold price surprises are relevant instruments for uncertainty shocks based on F-tests and Granger-causality tests with the VXO and the macro uncertainty measure constructed in Jurado et al. (2015). Ludvigson et al. (forthcoming) also use gold price changes as a proxy variable for global uncertainty shocks; Engel & Wu (2018) use the gold price as a proxy for uncertainty. Regarding the exogeneity condition $E[p_t^{\epsilon,u}\boldsymbol{\epsilon}_t^o] = 0$ in Equation (10b), Piffer & Podstawski (2018) document that gold price surprises are uncorrelated with a range of non-uncertainty shocks.[7,8]

Second, in the relevance condition in Equation (10a) we assume that US monetary policy shocks drive the Federal Funds Futures surprises on FOMC meeting days, $E[p_t^{\epsilon,mp}\epsilon_t^{mp}] \neq 0$. This is the standard instrument relevance assumption maintained in the literature (Gertler & Karadi 2015; Caldara & Herbst 2019; Jarociński & Karadi 2020). Regarding the exogeneity condition $E[p_t^{\epsilon,mp}\boldsymbol{\epsilon}_t^o] = 0$ in Equation (10b), it seems plausible that in a narrow time window around FOMC meetings monetary policy shocks are the only systematic drivers of Federal Funds Futures surprises, especially when these are cleansed from central bank information effects.

[7] The exogeneity condition for the gold price surprises might be questioned as on some of the dates also non-uncertainty shocks may have materialised. However, note that the events considered by Bloom (2009), Piffer & Podstawski (2018) as well as Bobasu et al. (2020) are very diverse, meaning that even if on each and every event it was not only a global uncertainty shock that materialised, the non-uncertainty shock is likely to have been of a different nature across events. For example, while the collapse of AIG may have been more a financial than a global uncertainty shock, the 9/11 attacks or the launch of Operation Desert Storm were arguably no financial shocks. Therefore, we believe it can be argued that the only structural shock that has been *systematically* related to gold price surprises across all dates selected by Bloom (2009), Piffer & Podstawski (2018) as well as Bobasu et al. (2020) are global uncertainty shocks.

[8] It should be clear that we do not include the VXO in the BPSVAR model as a measure of global uncertainty in order to identify the associated structural shocks. Instead, the identification of global uncertainty shocks rests on the assumptions on the relationship between structural shocks and proxy variables in Equation (10a), and are not tied to the specification of the vector of endogenous variables $\boldsymbol{y}_t$.

As discussed in Section 2, when multiple proxy variables are used to identify multiple structural shocks, the relevance and exogeneity conditions are not sufficient for point identification. In the BPSVAR framework additional restrictions can be imposed on $\boldsymbol{V}$. A natural idea is to impose that $\boldsymbol{V}$ is a diagonal matrix, implying that Federal Funds Futures surprises on FOMC meeting days are not driven by global uncertainty shocks and that gold price surprises on days with large exogenous increases in uncertainty are not driven by US monetary policy shocks. Technically, these additional restrictions imply an over-identified system, which cannot be handled by the estimation algorithm of Arias et al. (forthcoming). We therefore impose a *weaker* set of additional restrictions, namely only that Federal Funds Futures surprises on FOMC meeting days are not driven by global uncertainty shocks, $E[p_t^{\epsilon,mp}\epsilon_t^u] = 0$. Note that this assumption is implicitly maintained and crucial for the validity of much work in the literature. For example, if this assumption was not satisfied then the analyses of Gertler & Karadi (2015), Caldara & Herbst (2019) as well as Jarociński & Karadi (2020) would be invalid as the identified US monetary policy shocks would be contaminated by global uncertainty shocks. Moreover, recall that as in Jarociński & Karadi (2020) we cleanse the Federal Funds Futures surprises we use as proxy variable for the US monetary policy shock from central bank information shocks, which could be interpreted as uncertainty shocks as they allow agents to predict the future more accurately.

Our relevance conditions in Equation (10a) are also weak in the sense that they allow gold price surprises to be driven at least in part by US monetary policy shocks, that is we allow for $E[p_t^{\epsilon,u}\epsilon_t^{mp}] \neq 0$. Nevertheless, below we consider a robustness check in which we relax the assumption $E[p_t^{\epsilon,u}\epsilon_t^{\ell}] = 0$ for $\ell \neq u$, by replacing the corresponding zero restrictions in Equations (10a) and (10b) with restrictions on their relative magnitude.

Finally, it is worthwhile to point out that when two proxy variables are used to identify two structural shocks, a single additional zero restriction on $\boldsymbol{V}$ is sufficient for point-identification (Giacomini et al. forthcoming).[9]

Previous work has largely relied on the use of recursiveness assumptions to identify uncertainty shocks in SVAR models (Bloom 2009; Jurado et al. 2015; Baker et al. 2016; Basu & Bundick 2017). In particular, it is typically assumed that only uncertainty shocks have a contemporaneous effect on the uncertainty indicator in the VAR model. However, quite some evidence suggests that also other structural shocks may impact uncertainty contemporaneously. For example, US monetary policy shocks have been found to contemporaneously impact global uncertainty (Bekaert et al. 2013; Rey 2016; Miranda-Agrippino & Rey 2020); in fact, without constraining the contemporaneous responses we also find this in our application. Identifying global uncertainty shocks using proxy variables in the BPSVAR framework allows us to avoid imposing such recursiveness assumption.[10]

[9]This is appealing also because under set-identification credible sets are wider and results may depend on the choice of the prior distribution for the construction of the rotation matrices in the estimation (Baumeister & Hamilton 2015).

[10]Other approaches to overcome the limitations of recursive identification in the context of global uncertainty shocks include Carriero et al. (2019) in a Bayesian stochastic volatility VAR model, Alessandri et al. (2020) who follow Gazzani & Vicondoa (2020) and use a bridge-proxy SVAR model that imposes recursiveness only at a higher frequency, and Redl (2020) who uses narrative restrictions based on close election outcomes.

3.4 Priors

We use flat priors for the VAR parameters. We follow Caldara & Herbst (2019) as well as Arias et al. (forthcoming) and impose a 'relevance threshold' to express our prior belief that the proxy variables are relevant instruments. In particular, we require that at least a share $\gamma = 0.1$ of the variance of the proxy variables is accounted for by the US monetary policy and global uncertainty shocks, respectively; this is weaker than the relevance threshold of $\gamma = 0.2$ used by Arias et al. (forthcoming), and—although hard to compare conceptually—lies below the 'high-relevance' prior of Caldara & Herbst (2019).

4 Results

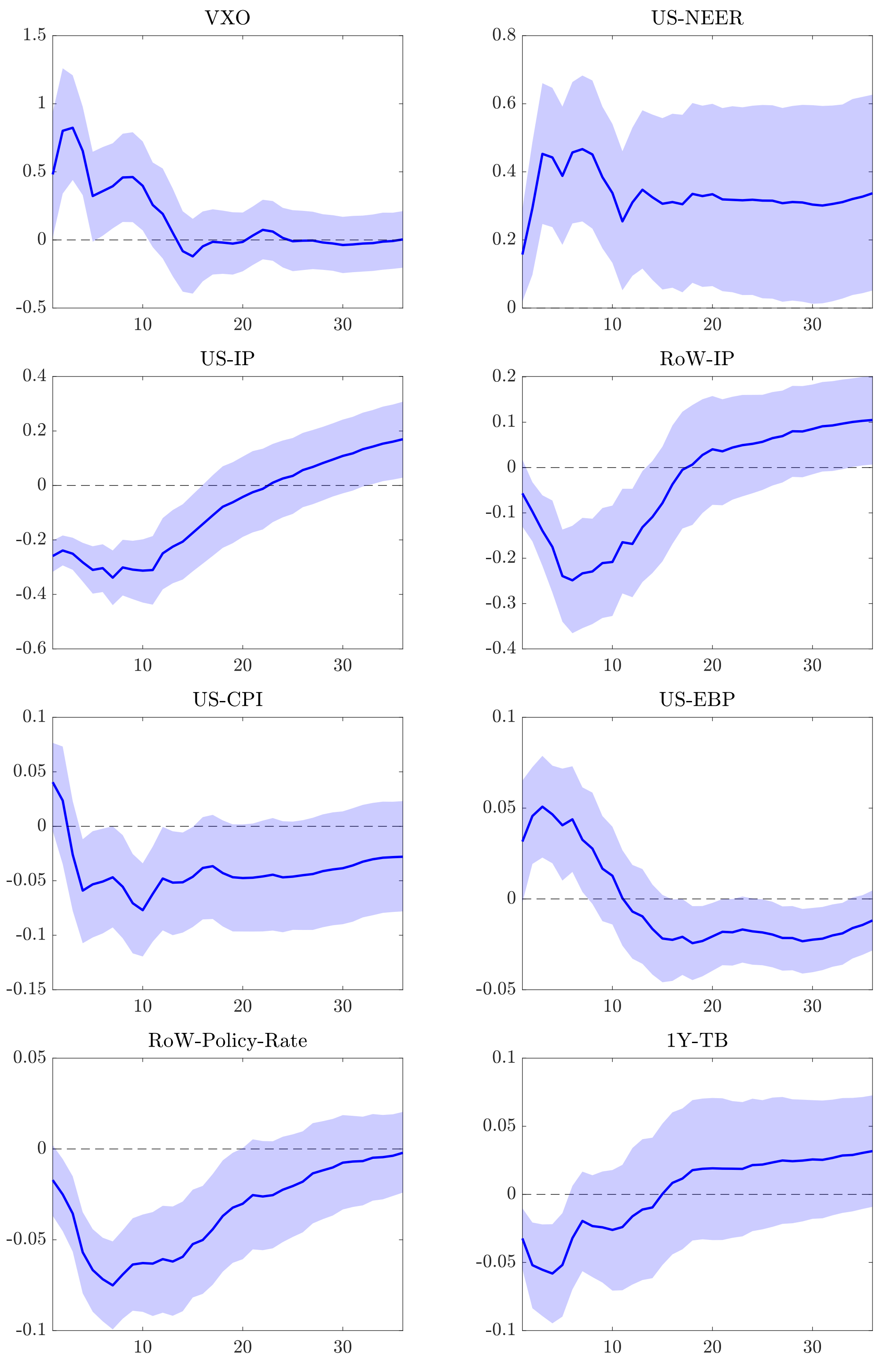
We first present results for the effects of global uncertainty shocks and then turn to a counterfactual analysis in order to explore the role of the dollar exchange rate. Finally, we explore the potential of US monetary policy to alleviate the effects of global uncertainty shocks. All results we report are based on the estimation of the BPSVAR model in which we identify a global uncertainty and a US monetary policy shock jointly.

4.1 The effects of global uncertainty shocks

Figure 2 displays the impulse responses to a one-standard deviation shock to global uncertainty. The VXO rises on impact and reaches its peak of about 1 point above baseline after three months. The response of the VXO is persistent: It takes about one year for the effect on the VXO to disappear. The dollar exchange rate exhibits some (delayed) overshooting, appreciating on impact by about 0.2% and for several months thereafter up to 0.4%. The appreciation is partly reversed after about a year, but not fully—the dollar remains appreciated relative to the pre-shock level by about 0.3%. The responses of US and RoW industrial production are similar. In both cases, there is a sharp contraction with a trough of about 0.3% reached after six months. This pattern has been documented as a distinct feature in the adjustment to uncertainty shocks (Bloom 2009). US consumer prices drop persistently, by about 0.05%. The external bond premium increases sharply on impact by about five basis points, and remains above baseline for almost one year. Interest rates fall both in the US and in RoW by up to eight basis points, reflecting that monetary policy accommodates the contractionary effect of the global uncertainty shock. In sum, we find that a global uncertainty shock has large effects on the world economy, with a fairly symmetric impact on the US and RoW. A distinctive asymmetry is the strong and persistent dollar appreciation.

Figure 3 presents the responses of other currencies' exchange rates to the global uncertainty shock obtained from augmenting $\boldsymbol{y}_t$ in the BPSVAR model with an additional variable one at a time. The top row presents responses for the Japanese yen and the Swiss franc, widely documented to be safe haven currencies in times of elevated uncertainty and risk-off episodes (Ranaldo & Söderlind 2010; De Bock & de Carvalho Filho 2015). The bottom panel presents the responses of the euro and the British pound as 'placebo' currencies. Consistent with the findings for the dollar, both the

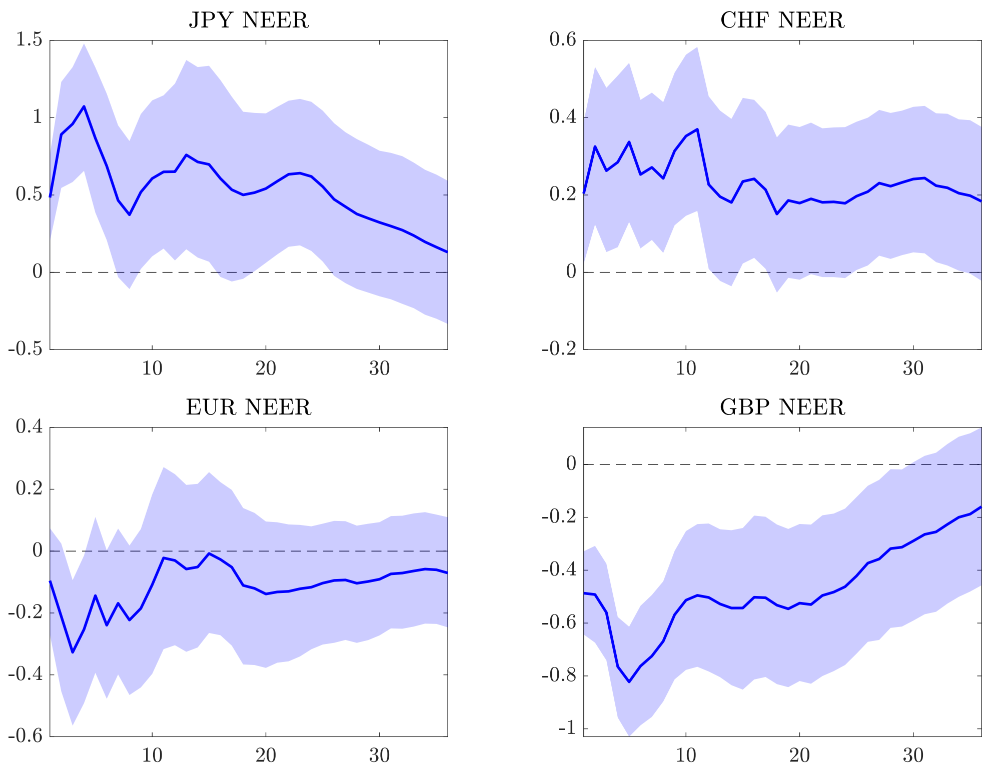
Figure 2: Impulse responses to global uncertainty shock



Note: The figure presents the impulse responses to a one-standard deviation global uncertainty shock. The blue solid line represents the point-wise posterior mean and the shaded areas 68% equal-tailed, point-wise credible sets. The VXO is depicted in levels, the dollar NEER, US and RoW industrial production, US consumer prices in logs, and the excess bond premium, the RoW policy as well as the US one-year Treasury Bill rates in percent.

Japanese yen and the Swiss franc appreciate in response to a global uncertainty shock. In contrast, the euro and the British pound depreciate.

Figure 3: Responses of the Japanese yen, Swiss franc, euro and pound sterling exchange rates to a global uncertainty shock



Note: See the notes to Figure 2. The responses are obtained from estimations of the BPSVAR model in which $\boldsymbol{y}_t$ is augmented with an additional variable one at a time

Before we explore the role of the dollar exchange rate in more detail, we consider two exercises to dispel concerns that our identified uncertainty shocks are contaminated by global demand shocks. First, we identify a global demand shock in addition to the global uncertainty and US monetary policy shocks using standard sign restrictions; importantly, we leave the response of the dollar exchange rate unconstrained[11]. Figure A.2 documents that the responses of all variables to the global demand shock are qualitatively consistent with those from the global uncertainty shock. The only exception is the dollar exchange rate, which appreciates much less in response to a global demand shock; on impact, there is even no appreciation at all. Figure A.3 documents analogous findings for the Japanese yen, the Swiss franc, the euro and the British pound. These results suggest that our baseline findings are unlikely to reflect the effects of global demand rather than uncertainty shocks.

Second, we relax the assumption in Equation (10b) that the proxy variable we use to identify the

[11]We impose the contemporaneous restrictions that the global demand shock reduces industrial production in the US and the rest of the world, that it reduces US consumer prices as well as policy rates in the US and RoW, and that it raises the excess bond premium.

global uncertainty shock—the gold price surprises—is uncorrelated with the non-uncertainty and non-US monetary policy structural shocks in $\boldsymbol{\epsilon}_t^o$. In particular, as in Ludvigson et al. (forthcoming) we instead allow the structural shocks in $\boldsymbol{\epsilon}_t^o$ to be correlated with the gold price surprises and only assume that the correlation is strongest with the global uncertainty shock, that is $|E[p_t^{\epsilon,u}\epsilon_t^u]| > |E[p_t^{\epsilon,u}\epsilon_t^\ell]|$ for $\ell \neq u$. Figure A.4 documents that the responses to the global demand shock we obtain from this alternative identification are very similar to those from the baseline.

Finally, note that the estimated effects of a global uncertainty shock in Figure 2 are qualitatively inconsistent with those of news shocks as estimated for example in Piffer & Podstawski (2018). In particular, Piffer & Podstawski (2018) jointly set-identify uncertainty and news shocks, finding that while uncertainty shocks are followed by a US monetary policy easing and a decline in US inflation, the opposite materialises following a news shock. Hence, our identified global uncertainty shocks are unlikely to be contaminated to a noteworthy degree by news shocks.

4.2 The role of the dollar

While in theory expenditure switching in response to dollar appreciation and wealth transfers in the context of the exorbitant duty imply expansionary output effects in RoW (Obstfeld & Rogoff 1996; Gourinchas & Rey 2007; Gourinchas et al. 2010), the financial channel of exchange rates implies contractionary effects (Bruno & Shin 2015).

The *trade channel* centers on expenditure switching that is triggered by dollar appreciation and which shifts demand away from goods produced in the US towards goods produced in RoW (Obstfeld & Rogoff 1996). In other words, under the trade channel dollar appreciation improves RoW net exports and—as a mirror image—worsens US net exports. The adjustments in net export patterns are—*ceteris paribus*—expansionary for RoW, and hence *dampen* contractionary effects of global uncertainty shocks.

The *insurance channel* rests on wealth transfers from the US to RoW that materialise as a result of valuation effects (Gourinchas & Rey 2007; Gourinchas et al. 2010). A key ingredient is the stylized fact that the US external balance sheet is long in risky, foreign-currency denominated foreign equity, and short in dollar-denominated US debt securities—US Treasuries to a large extent. In this setting, a global uncertainty shock that appreciates the dollar exchange rate, boosts prices of safe US assets and depresses global equity prices entails a deterioration in the US net foreign asset position through negative price and exchange rate valuation effects, and a commensurate improvement in the RoW net foreign asset position. Hence, the dollar appreciation *dampens* the contractionary effects of global uncertainty shocks through the insurance channel.

The *financial channel* represents the tightening in global financial conditions triggered by dollar appreciation when globally active banks operate under value-at-risk constraints and intermediate dollar funding from the US to RoW (Bruno & Shin 2015). In this setting, a global uncertainty shock that appreciates the dollar exchange rate weakens local borrowers' balance sheets facing currency mismatches, increases their credit risk and eventually reduces lending capacity of globally active banks. Hence, the dollar appreciation *amplifies* the contractionary effects of global uncertainty shocks through the financial channel.

The contribution of the dollar appreciation to the overall output effects of the global uncertainty shock is hence ambiguous in theory. We explore a counterfactual exercise to assess empirically whether the exchange rate effects through the trade and insurance channel or the financial channel dominate when the dollar appreciates in the face of a global uncertainty shock. We first provide an intuitive discussion of counterfactual analysis in VAR models, and then lay out in more detail the approach we adopt.

4.2.1 Minimum relative entropy counterfactuals

In the existing empirical literature MRE is used to incorporate restrictions derived from economic theory in order to improve a forecast. For example, Robertson et al. (2005) improve their forecasts of the Federal Funds rate, US inflation and the output gap by imposing the constraint that the mean three-year-ahead inflation forecast must equal 2.5% through MRE.[12] Similar in spirit, as in Breitenlechner et al. (2020) we use MRE to generate a counterfactual conditional forecast based on our baseline conditional forecast that represents the impulse responses to a global uncertainty shock.

Conceive of an impulse response as a conditional forecast $\widetilde{\boldsymbol{y}}_{T+1,T+h} \equiv [\widetilde{\boldsymbol{y}}'_{T+1}, \widetilde{\boldsymbol{y}}'_{T+2}, \ldots, \widetilde{\boldsymbol{y}}'_{T+h}]'$ over periods $T+1, T+2, \ldots, T+h$ under which the structural shocks $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \equiv [\widetilde{\boldsymbol{\epsilon}}'_{T+1}, \widetilde{\boldsymbol{\epsilon}}'_{T+2}, \ldots, \widetilde{\boldsymbol{\epsilon}}'_{T+h}]'$ assume the values $\widetilde{\epsilon}^u_{T+1} = 1$, $\widetilde{\epsilon}^u_{T+s} = 0$ for $s = 2, 3, \ldots, h$ and $\widetilde{\epsilon}^\ell_{T+s} = 0$ for $s = 1, 2, \ldots, h$ and $\ell \neq u$. Furthermore, consider the posterior belief about the effects of a US monetary policy shock based on the actual data

$$f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T}, \mathcal{I}_a, \widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}) \propto p(\boldsymbol{\psi}) \times \ell(\boldsymbol{y}_{1,T}|\boldsymbol{\psi}, \mathcal{I}_a) \times \nu, \tag{11}$$

where $p(\cdot)$ is the prior about the structural VAR parameters $\boldsymbol{\psi}$, $\mathcal{I}_a$ our identifying assumptions, and ν the volume element of the mapping from the posterior distribution of $\boldsymbol{\psi}$ to the posterior distribution of the conditional forecast $\widetilde{\boldsymbol{y}}_{T+h}$; the mean of f is represented by the blue solid lines in Figure 2. MRE determines the posterior beliefs about the effects of a global uncertainty shock in a counterfactual VAR model by

$$\begin{gathered} Min_{\boldsymbol{\psi}} \ \mathcal{D}(f^*||f) \quad s.t. \\ \int f^*(\widetilde{\boldsymbol{y}})\widetilde{y}^{\$} d\widetilde{\boldsymbol{y}} = E(\widetilde{y}^{\$}) = 0, \ \ \int f^*(\widetilde{\boldsymbol{y}}) d\widetilde{\boldsymbol{y}} = 1, \ \ f^*(\widetilde{\boldsymbol{y}}) \geq 0, \end{gathered} \tag{12}$$

where $\mathcal{D}(\cdot)$ is the Kullback-Leibler divergence—the 'relative entropy'—between the counterfactual and actual posterior beliefs (we drop the subscripts in $\widetilde{y}^{\$}_{T+h}/\widetilde{\boldsymbol{y}}_{T+h}$ in Equation (12) for simplicity). In general, there is an infinite number of counterfactual VAR models about which given the data $\boldsymbol{y}_{1,T}$ we could construct beliefs f^* that satisfy the constraint $E(\widetilde{y}^{\$}_{T+h}) = 0$. The MRE approach in Equation (12) disciplines the choice of the counterfactual VAR model in an arguably plausible way by requiring that it is associated with beliefs f^* that are minimally different from the baseline posterior beliefs f in an information-theoretic sense. Intuitively and roughly speaking, MRE determines that

[12] See Cogley et al. (2005) and Giacomini & Ragusa (2014) for similar applications.

counterfactual VAR model which is minimally different from the actual VAR model but which features that the dollar exchange rate is unresponsive to a global uncertainty shock.[13]

It turns out that in order to determine the posterior beliefs f^* about the effects of a global uncertainty shock in Equation (12) MRE optimally updates the baseline posterior beliefs f by incorporating the information represented by the constraint that the dollar exchange rate is expected to be unresponsive in the counterfactual VAR model according to

$$\begin{aligned} f^*\left(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T},\mathcal{I}_a,\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h},E(\widetilde{y}^{\$}_{T+h})=0\right) \propto \\ f(\widetilde{\boldsymbol{y}}_{T+h}|\boldsymbol{y}_{1,T},\mathcal{I}_a,\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}) \times \tau\left(\widetilde{y}^{\$}_{T+h}(\boldsymbol{\psi})\right), \end{aligned} \tag{13}$$

where τ is a 'tilt' function (see Robertson et al. 2005).[14] Intuitively, τ down-weights the baseline posterior for VAR parameter values $\boldsymbol{\psi}$ that are associated with large deviations from the constraint that the dollar exchange rate shall be unresponsive. In practice, Robertson et al. (2005) as well as Giacomini & Ragusa (2014) show that MRE boils down to adjusting the weights of the draws of the approximated baseline posterior distribution. Once the counterfactual weights are obtained, importance sampling techniques can be used to estimate the mean and percentiles of the counterfactual posterior distribution.[15]

4.2.2 Results from MRE no-appreciation counterfactuals

Figure 4 shows the result of the MRE approach for the role of the dollar in the transmission of global uncertainty shocks. In each panel, the blue solid line represents the impulse responses from the baseline, and the circled red line represents the impulse responses from the MRE counterfactual in which the dollar exchange rate does not respond to the global uncertainty shock.

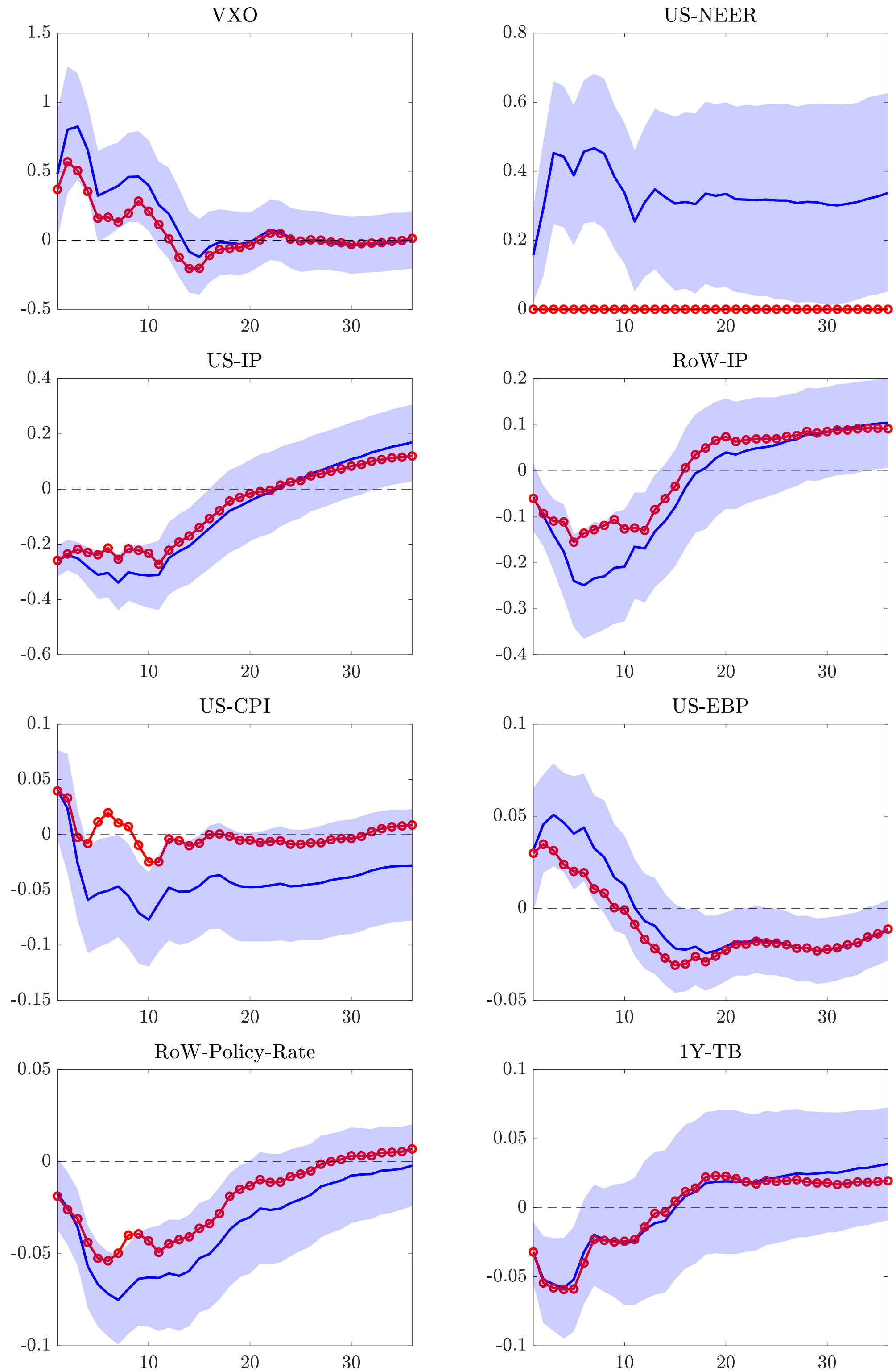
In the counterfactual the VXO and the excess bond premium rise by less in response to a global uncertainty shock than in the baseline. US and RoW industrial production, US consumer prices as well as US and RoW short-term interest rates fall by less than in the baseline. The MRE counterfactual thus suggests that dollar appreciation amplifies the effects of global uncertainty

[13] Brute force alternatives for carrying out counterfactual analysis in VAR models are to set to zero autoregressive parameters after or before estimation (see, for example, Carriere-Swallow & Cespedes 2013; Dees & Galesi 2019; Vicondoa 2019; Degasperi et al. 2020; Redl 2020). However, setting to zero VAR coefficients in general either violates the Lucas critique or implies a mis-specified empirical model and induces biased estimates (see Georgiadis 2017).

[14] Beliefs can be updated not only based on data but based on any form of new information. Optimally updating beliefs based on data refers to Bayes' rule. In case of other information—such as the constraint that real activity spillovers in the counterfactual VAR model are expected to be nil—it can be shown that MRE updating as in Equation (13) is optimal in an axiomatic sense (see for example Shore & Johnson 1980, 1981). Also note that there is nothing dubious about using actual data in MRE to form beliefs about a counterfactual VAR model. For example Giffin (2008, pp. 25-26) writes: "The distribution that we get in the end is based on our information, not what is or is not true".

[15] Importance sampling is only feasible and efficient if the baseline density—in our case the posterior distribution of the impulse responses—spans the target density. As shown in Arias et al. (2018) the posterior of the impulse responses follows a Generalized-Normal distribution, which has infinite support in theory. Hence, any counterfactual posterior distribution of conditional forecasts can be obtained using MRE updating in theory. However, in practice when the posterior distribution is approximated by a finite number of draws and when the target density is very different from the baseline density, importance sampling might perform poorly. In this case, other sampling techniques can be used, for instance the one-block tailored Metropolis–Hastings algorithm of Chib et al. (2018).

Figure 4: Baseline and MRE-based counterfactual responses to a global uncertainty shock



Note: See the notes to Figure 2. The red dotted lines depict point-wise means of the counterfactual posterior distribution obtained from the MRE approach.

shocks on real activity in the US and RoW, adds downward pressure on US consumer prices, and elicits a stronger loosening of monetary policy in RoW. The counterfactual also suggests that dollar appreciation amplifies the rise in uncertainty as measured by the VXO as well as the tightening in financial conditions reflected in the excess bond premium. Finally, because the counterfactual suggests that dollar appreciation overall amplifies the contractionary effects of global uncertainty shocks, it also implies that the financial channel is more powerful than the trade and insurance channels.

4.2.3 Inspecting the mechanisms

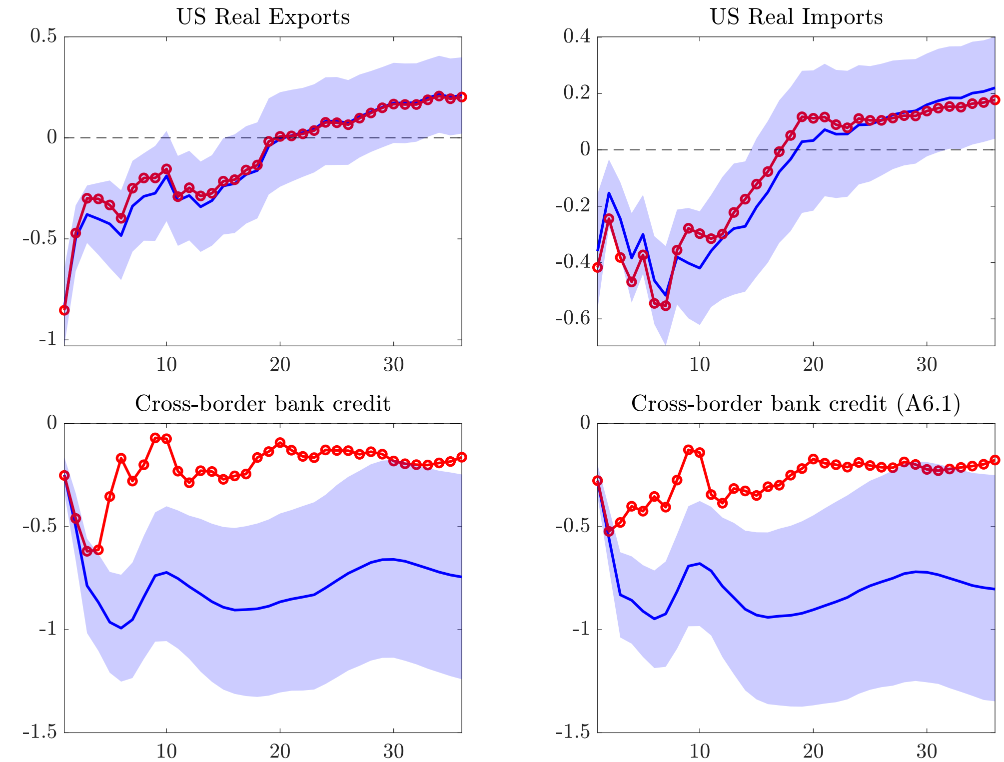
To corroborate this conclusion, we estimate the effects of global uncertainty shocks on the variables reflecting the transmission of dollar appreciation through the financial and the trade channel. In particular, we augment the BPSVAR model one at a time with cross-border bank credit to non-US borrowers, US real exports and imports.

As in Figure 4, the solid blue lines in Figure 5 present the baseline responses of US exports and imports as well as cross-border bank credit to non-US borrowers to a global uncertainty shock. For cross-border bank credit, the left-hand side panel depicts results for data based on the nationality principle applied to the reporting banks and the left-hand side based on the residency principle.[16] The results indicate that in response to a global uncertainty shock cross-border bank credit to non-US borrowers, US exports and US imports all decline. The drop in US exports is more pronounced than the drop in US imports, implying an expansionary contribution of US net exports to RoW real activity. While the latter is consistent with expenditure switching in the face of dollar appreciation, it could also be driven entirely by demand effects given that real activity slows down in response to the global uncertainty shock both in the US and RoW. Coincidentally, our findings are consistent with dominant-currency pricing (DCP; Gopinath et al. 2020): When both US import and export prices are sticky in US dollar terms in the short run, dollar appreciation only elicits expenditure switching on US exports; as both the US and the RoW experience similar negative demand effects, DCP implies a stronger short-term drop in US exports than imports.

The circled red lines in Figure 4 present the responses for the MRE counterfactual. The results corroborate our main finding that the financial channel dominates the trade channel. In particular, when dollar appreciation is absent, cross-border bank credit to non-US borrowers declines by much less than in the baseline, implying considerably less pronounced tightening of financial conditions in

[16] In the left-hand side panel we consider the data from Table A7 from the BIS Locational Banking Statistics on "External liabilities to all sectors of all reporting banks" less "External liabilities to all sectors of banks owned by US nationals" (see Table A.1 for variable definitions/descriptions and Figure A.6 for the evolution over time). This is the same data Bruno & Shin (2015) use in their analysis. The advantage of this data is that it is based on the nationality principle, meaning that distortions introduced through financial centers are reduced. The disadvantage is that the data only reflect information on the liabilities of globally active banks in the BIS reporting countries, which ranged between 24 in the 1990s and 48 at the end of our sample period (BIS 2020), potentially omitting some important emerging market and small open economies. At the same time, it is estimated that the coverage of the cross-border claims of all banks worldwide even in 1990s amounted to about 90%. In the right-hand side panel the data on cross-border bank credit are taken from Table A6.1 of the BIS Locational Banking Statistics based on residency principle, and the variable is calculated as "Banks' external claims on all sectors in all countries" less "Banks' external claims on all sectors in the US".

Figure 5: Baseline and MRE-based counterfactual responses of US trade and cross-border bank credit to a global uncertainty shock



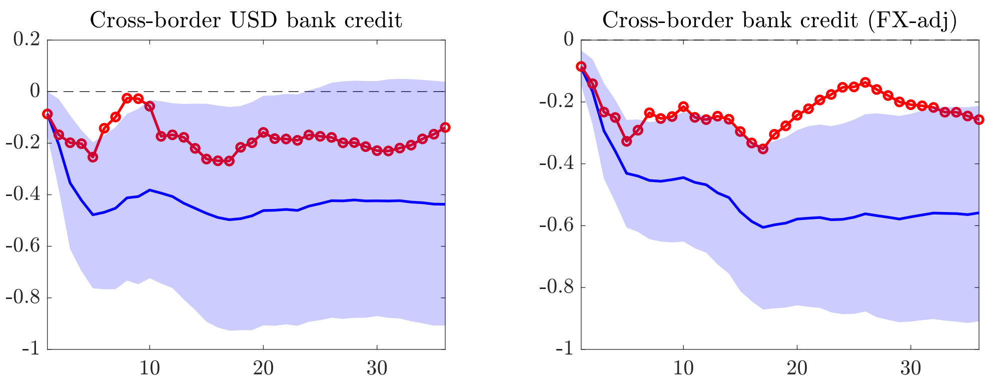
Note: See the notes to Figure 4. In the lower left-hand side panel the data on cross-border bank credit are taken from the BIS Locational Banking Statistics Table A7 based on nationality principle, and the variable is calculated as "External liabilities to all sectors of all reporting banks" less "External liabilities to all sectors of banks owned by US nationals". In the lower right-hand side panel the data on cross-border bank credit are taken from Table A6.1 based on residency principle, and the variable is calculated as "Banks' external claims on all sectors in all countries" less "Banks' external claims on all sectors in the US".

RoW. Regarding US trade, the absence of dollar appreciation entails a weaker drop in US exports and a stronger drop in US imports but the difference to the baseline is very small.

Figure A.5 documents that the results are very similar if we do not add US exports and imports as well as cross-border bank credit to the BPSVAR model one at a time but instead include them simultaneously. In order to account for the greater dimensionality of the specification, in this case we estimate the BPSVAR model with informative Minnesota-type priors and optimal hyperpriors/prior tightness as in Giannone et al. (2015).

A possible objection to the finding of a much weaker drop in cross-border bank credit in the counterfactual is that this might be a statistical artifact due to the recording of non-US dollar denominated credit flows in US dollar. In particular, when the dollar appreciates in response to a global uncertainty shock in the baseline, then any cross-border bank credit flows denominated in non-US dollar currencies imply a reduced dollar value. When the dollar exchange rate is unresponsive in the counterfactual, this mechanical valuation effect is absent, reducing the drop in cross-border bank credit in response to the global uncertainty shock. Figure 6 presents the results for the effects of global uncertainty shocks for two alternative variables for cross-border bank credit, namely only US dollar denominated cross-border bank credit—which accounts for about half of total cross-border bank credit (see Figure A.7)—and cross-border bank credit adjusted for exchange rate movements. While both the baseline drop in cross-border bank credit and the reduction in the drop in the counterfactual is indeed smaller for these two alternative cross-border bank credit variables than in Figure 5, both the baseline drop and the reduction of the drop in the counterfactual remain substantial.

Figure 6: Baseline and MRE-based counterfactual responses of alternative cross-border bank credit variables



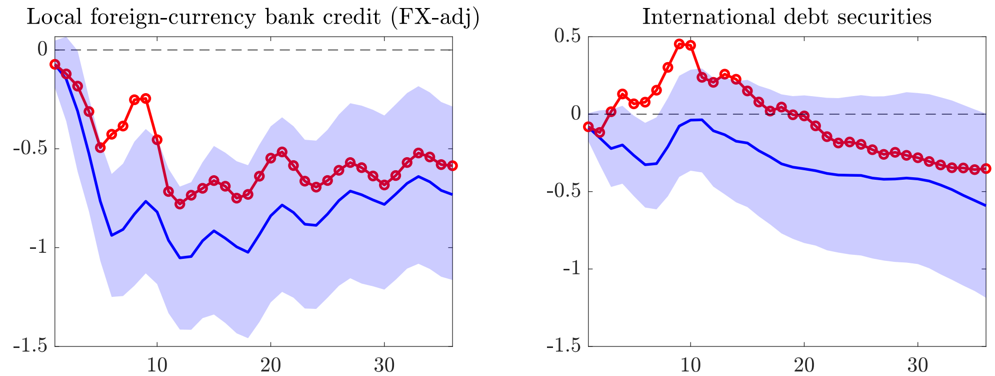
Note: See the notes to Figure 4. The left-hand side panel depicts the responses for US dollar instead of total cross-border bank credit and the right-hand side panel for the exchange rate-adjusted total cross-border bank credit.

In sum, US net exports fall by less in response to a global uncertainty shock when dollar appreciation is absent in the counterfactual. All else equal, the weaker drop in US net exports implies a harsher slowdown of economic activity in the rest of the world, while a weaker drop in cross-border bank credit implies a weaker slowdown. As we find that economic activity in RoW contracts considerably less in the counterfactual, the financial channel must dominate the trade channel. In other words, the main contribution of dollar appreciation in transmitting a global uncertainty shock to RoW appears to be through the financial channel of exchange rates.

The critical role of the effect of global uncertainty shocks and the role of the dollar in their transmission is reflected in further credit variables. In particular, the left-hand side panel in Figure 7 documents that also local foreign-currency credit by globally active banks drops substantially in response to a global uncertainty shock, and that this drop is reduced substantially—especially over a horizon of up to one year—when dollar appreciation is precluded.[17] The right-hand side panel shows that qualitatively similar findings are obtained for international debt securities issued outside the jurisdiction in which the issuer resides.

[17]Local credit are claims on counterparties located in the same country as the banking group's entity that books the claim. In contrast, cross-border credit reflect claims on counterparties located outside the country in which the entity that books the position is located (BIS 2019).

Figure 7: Baseline and MRE counterfactual responses of local foreign-currency bank credit and international debt securities



Note: See the notes to Figure 4. The left-hand side panel depicts the responses of local foreign-currency credit obtained from the BIS Locational Banking Statistics defined as "Local claims in foreign currencies of banks owned by the world" less "Local claims in foreign currencies of banks owned by the US". The right-hand side panel depicts the responses of international debt securities issued in jurisdictions other than the US obtained from the BIS International Debt Securities database.

Our results extend earlier work on the role of financial factors for the transmission of global uncertainty shocks. For example, Carriere-Swallow & Cespedes (2013) consider uncertainty shocks given by changes in the VIX that exceed some pre-specified threshold in small-open economy VAR models for 20 EMEs and 20 AEs, finding that EMEs suffer deeper and more prolonged contractions than AEs and that this is largely due to credit constraints. Carriere-Swallow & Cespedes (2013) do not explore the role of the dollar exchange rate for the cross-border transmission of US uncertainty shocks. Liu et al. (2017) estimate factor-augmented VAR models and find that innovations to the dollar effective exchange rate ordered last are followed by a slowdown in real activity in South Korea, and argue this is due to negative US demand effects outweighing expenditure switching. However, Liu et al. (2017) do not consider the role of the financial channel as a competing explanation. Interestingly, Liu et al. (2017) find that dollar appreciation hardly affects China and Japan, both of which are arguably not dependent on cross-border bank credit; moreover, tit may be that the Japanese yen does not depreciate against the dollar when uncertainty rises due to its safe-haven property. Shousha (2019) estimates panel VAR models for EMEs and finds that innovations to the dollar NEER ordered last among US variables are followed by contractions in output abroad, which are deeper for countries with greater balance sheet exposures in terms of the share of credit to the non-financial private sector that is denominated in dollar; but it is not clear to what extent the shocks considered reflect global uncertainty or other structural shocks. Cesa-Bianchi et al. (2018) estimate the effects of international credit supply shocks—*inter alia* reflecting risk appetite shocks—identified by innovations to US broker-dealer leverage ordered first in a mean-group panel VAR model and find that they cause dollar appreciation, a decline in consumption and cross-border banking claims; but they do not explore the contribution of dollar appreciation for the overall contraction. Epstein et al.

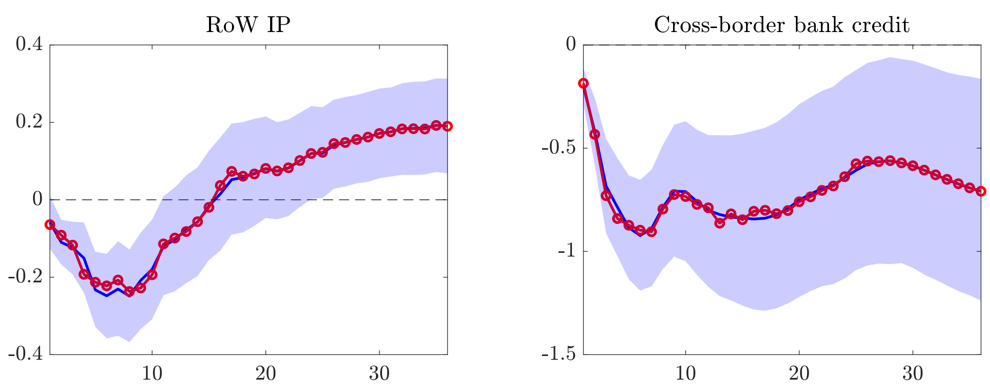
(2019) identify global financial risk shocks as innovations to the US BAA spread ordered last in small open economy panel VAR models and find that bank credit plays a key role at least for advanced economies; but they again do not link the behaviour of bank credit to the dollar appreciation.

We next explore whether our focus on the dollar is also justified by a uniquely important role for the transmission of global uncertainty shocks. We consider first the dollar exchange rate itself contrasting it against other safe haven currencies, and then dollar relative to non-dollar cross-border credit.

4.2.4 The unique role of the dollar exchange rate

The left-hand side panel in Figure 8 presents the response of RoW industrial production for the MRE counterfactual in which the Japanese yen is unresponsive to the global uncertainty shock while the dollar exchange rate is assumed to respond as in the baseline; the right-hand side panel shows the corresponding counterfactual responses of cross-border bank credit. The results suggest that precluding the appreciation of the Japanese yen in a response to a global uncertainty shock is not associated with a weaker contraction of the world economy and cross-border bank credit; the results are very similar for the Swiss franc. Hence, unlike the dollar exchange rate the behaviour of other safe haven currencies is inconsequential for the effects of global uncertainty shocks. Plausibly, the reason for this finding is that cross-border bank credit in other safe haven currencies is minuscule (see Figure A.7). The only currency which accounts for a quantitatively similar share of cross-border bank credit as the dollar is the euro, whose exchange rate we document in Figure 3 has not appreciated in response to global uncertainty shocks. We turn to the comparison of dollar and euro cross-border bank credit next.

Figure 8: Baseline and MRE-based counterfactual responses of when Japanese yen instead of US dollar is unresponsive to a global uncertainty shock



Note: See the notes to Figure 4. The red dotted lines depict the responses of rest-of-the-world industrial production (left-hand side panel) and cross-border bank credit (right-hand side panel) in the counterfactual in which the Japanese yen is constrained to not respond to the global uncertainty shock.

4.2.5 The unique role of dollar cross-border bank credit

Bruno & Shin (2015) are concerned with the relationship between of dollar appreciation and the overall lending of globally active banks, but they do not distinguish between cross-border bank credit in different currencies. Ivashina et al. (2015) present a model in which globally active banks cut dollar lending more than euro lending in response to a shock to their credit quality. In particular, because globally active banks rely on unsecured dollar funding through wholesale markets in the US while raising euro through insured retail deposits, a credit quality shock leads to a greater drop in dollar funding. In principle, banks could borrow in euros and swap into dollars to make up for the dollar funding shortfall, but this is precluded by violations of covered interest parity that arise when there is limited capital to take the other side of the swap trade. As a result, the credit quality shock causes cuts in dollar but not in euro lending.

Parsed into the context of our paper, the exchange rate responses to a global uncertainty shock entail that borrowers with dollar but not euro mismatches on their balance sheets become more risky. The prediction that emerges is that dollar cross-border bank credit should *ceteris paribus* be impacted more strongly by the global uncertainty shock than non-dollar cross-border bank credit.[18] In line with the analysis of Ivashina et al. (2015), Avdjiev, Du, et al. (2019) document a 'triangular' relationship in that a (i) stronger dollar goes hand in hand with (ii) larger deviations from covered interest parity and (iii) contractions of dollar cross-border bank credit.

To test this additional prediction about the special role of the dollar under the financial channel we re-do the estimation and the counterfactual analysis separately for dollar and non-dollar cross-border bank credit.[19]

Notice that contrasting dollar and non-dollar cross-border bank credit responses to a global uncertainty shock also allows us to disentangle the role of cross-border bank credit demand and supply. In particular, while we expect dollar and non-dollar cross-border bank credit to drop—*ceteris paribus*—equally in response to a contractionary global uncertainty shock due to credit demand, in the environment spelled out in Ivashina et al. (2015) the financial channel predicts dollar cross-border

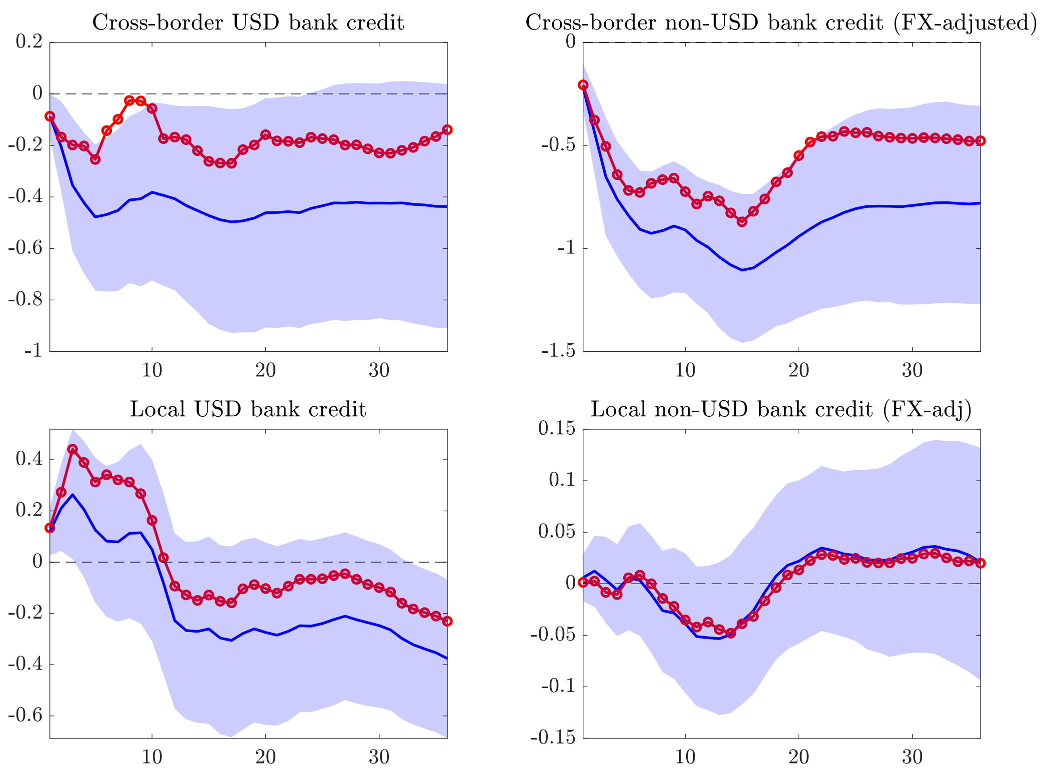
[18] Takats & Temesvary (2020) study a 'currency dimension' of the international bank lending channel and find that changes in monetary policy in the US, the euro area and Japan affect global cross-border bank lending denominated in dollar, euro and yen, respectively; for example, changes in US monetary policy affect cross-border dollar lending of a UK bank to a borrower in Malaysia. The finding is rationalised as higher liquidity in a currency that results from monetary policy easing by the issuer central bank raises funding available in that currency to globally active banks, which then translates into increased cross-border loan supply in that currency. Avdjiev et al. (2018) obtain similar findings. The mechanism we explore is different in the sense that it is not linked to the effect of monetary policy on funding costs.

[19] Non-dollar cross-border bank credit also includes yen-denominated flows. Because the yen exchange rate also appreciates in response to a global uncertainty shock (see Figure 3), we may expect similar mechanisms to play out for cross-border yen and dollar credit. However, recall that cross-border yen credit is quantitatively minuscule relative to dollar and euro credit (see Figure A.7), and hence at best we expect lumping yen with euro cross-border bank credit to make it more difficult to obtain results consistent with the prediction that dollar cross-border bank credit is special. Moreover, in the mechanism highlighted by Ivashina et al. (2015) differences in the responses of dollar and non-dollar cross-border credit result from differences in the sensitivity of funding to shocks to the riskiness of borrowers. While dollar funding originates from US wholesale markets, yen funding raised from Japanese deposits should arguably be more resilient. As a result, we expect cross-border bank credit in Japanese yen to be less sensitive to global uncertainty shocks than cross-border bank credit in dollars.

bank credit supply to drop more strongly than non-dollar cross-border bank credit supply in response to a global uncertainty shock. There may be differences across the responses of dollar and non-dollar cross-border bank credit supply due to other factors, such as systematic differences in the globally active banks extending the respective credit and their borrowers. An appealing feature of our counterfactual exercise is that these factors are constant in the baseline and the counterfactual. In that sense, the MRE counterfactual can be viewed as a diff-in-diff for assessing the potentially unique role of dollar cross-border bank credit in the transmission of global uncertainty shocks.

Our findings are consistent with this additional prediction from Ivashina et al. (2015). In particular, the panels in the top row Figure 9 documents that the reduction in the drop in response to a global uncertainty shock is greater for dollar than for non-dollar cross-border bank credit. That there is a reduction also in the decline of non-US dollar cross-border bank credit under the counterfactual is plausible, as it entails a weaker drop in global output, and hence weaker credit demand effects. The panels in the bottom row suggest results are similar when considering local instead of cross-border credit.

Figure 9: Baseline and MRE counterfactual responses of dollar and non-dollar cross-border and local credit



Note: See the notes to Figure 4.

4.3 Could US monetary policy alleviate the effects of global uncertainty shocks?

The appreciation and acute run for US dollar in times when global uncertainty spikes often triggers a response by the Federal Reserve. This was exemplified recently during the COVID-19 pandemic, when the Federal Reserve provided emergency liquidity to a number of countries through various facilities (see Cetorelli et al. 2020). Against the background of our findings, we now ask by how much the contractionary effects of the global uncertainty shock would be alleviated if the Federal Reserve were to prevent the appreciation of the dollar exchange rate. Since US monetary policy can impact the exchange rate through expected future interest rate differentials and risk premia, it is a natural question to ask to what extent the Federal Reserve could stabilise the global economy in the face of an uncertainty shock. We first briefly sketch the SSC methodology we apply for the purposes of this counterfactual exercise.

4.3.1 Structural shock counterfactuals

In SSCs a constrained conditional forecast is produced as additional shocks materialise along the forecast horizon. In the context of counterfactual analysis that evaluates the role of particular transmission channels, it is typically applied with a constraint on the set of shocks that may materialise (Kilian & Lewis 2011; Bachmann & Sims 2012; Wong 2015; Epstein et al. 2019). Intuitively, while MRE aims at exploring a counterfactual world while keeping future shocks at zero, SSCs explore a counterfactual scenario in the actual world in which a particular series of future shocks materialses. Antolin-Diaz et al. (2021, henceforth ADPRR) provide a coherent treatment of how to impose constrained paths on observables as conditional forecasts together with constraints on the set of driving shocks in a VAR model.[20] We provide a sketch of the framework here; a more detailed discussion in the context of counterfactual analysis is provided in Breitenlechner et al. (2020).

The values of the endogenous variables in a VAR model over a forecast horizon of h periods are given by

$$\boldsymbol{y}_{T+1,T+h} = \boldsymbol{b}_{T+1,T+h} + \boldsymbol{M}'\boldsymbol{\epsilon}_{T+1,T+h}, \tag{14}$$

where the $nh \times 1$ vector $\boldsymbol{b}_{T+1,T+h}$ represents the deterministic component of the forecast that is due to initial conditions and the autoregressive dynamics of the system, and the $nh \times nh$ matrix $\boldsymbol{M}'$ the effects of the structural shocks. The object of interest in SSCs is

$$\widetilde{\boldsymbol{y}}_{T+1,T+h} \sim N(\boldsymbol{\mu}_y, \boldsymbol{\Sigma}_y), \tag{15}$$

where the $nh \times 1$ vector $\widetilde{\boldsymbol{y}}_{T+1,T+h}$ contains all endogenous variables—i.e. both those whose paths are constrained and those whose paths are unconstrained. A conditional forecast in the framework of ADPRR involves

(i) ‘conditional-on-observables forecasting’, i.e. specifying paths for a subset of endogenous

[20] Leeper & Zha (2003) propose procedures for testing whether the (policy) shocks used in SSCs can be viewed as ‘modest policy interventions’ that are unlikely to induce agents to revise their beliefs about policy rules and hence do not raise the Lucas critique. Similarly, ADPRR propose the ‘q-divergence’ to judge how likely a counterfactual is.

variables that deviate from the unconditional forecast

(ii) 'conditional-on-shocks forecasting', i.e. specifying the subset of structural shocks that are allowed to deviate from their unconditional distribution to produce the path of the endogenous variables specified in (i)

Given Equation (14) 'conditional-on-observables forecasting' under (i) can be written as

$$\overline{\boldsymbol{C}}\widetilde{\boldsymbol{y}}_{T+1,T+h} = \overline{\boldsymbol{C}}\boldsymbol{b}_{T+1,T+h} + \overline{\boldsymbol{C}}\boldsymbol{M}'\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\overline{\boldsymbol{f}}_{T+1,T+h}, \overline{\boldsymbol{\Omega}}_f), \tag{16}$$

where $\overline{\boldsymbol{C}}$ is a $k_o \times nh$ selection matrix, the $k_o \times 1$ vector $\overline{\boldsymbol{f}}_{T+1,T+h}$ is the mean of the distribution of the endogenous variables that are constrained under the conditional forecast and the $k_o \times k_o$ matrix $\overline{\boldsymbol{\Omega}}_f$ reflects the associated uncertainty; in case of a SSCs $\overline{\boldsymbol{\Omega}}_f = \mathbf{0}$. Intuitively, the matrix $\overline{\boldsymbol{C}}$ selects the endogenous variables whose paths shall be constrained—i.e. the dollar exchange rate in the context of our paper. In turn, 'conditional-on-shocks forecasting' under (ii) can be written as

$$\boldsymbol{\Xi}\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h} \sim N(\boldsymbol{g}_{T+1,T+h}, \boldsymbol{\Omega}_g), \tag{17}$$

where $\boldsymbol{\Xi}$ is a $k_s \times nh$ selection matrix, the $k_s \times 1$ vector $\boldsymbol{g}_{T+1,T+h}$ the mean of the distribution of the shocks $\widetilde{\boldsymbol{\epsilon}}_{T+1,T+h}$ in the conditional forecast, and the $k_s \times k_s$ matrix $\boldsymbol{\Omega}_g$ reflects the associated uncertainty. ADPRR show how $\boldsymbol{\mu}_y$ and $\boldsymbol{\Sigma}_y$ in Equation (15) can be determined such that the constraints under (i) and (ii) are satisfied.

We again conceive as conditional forecasts both the baseline and the counterfactual impulse responses to a global uncertainty shock. In particular, our baseline is given by the forecast $\widetilde{\boldsymbol{y}}^{bl}_{T+1,T+h}$ conditional on a global uncertainty shock materialising in period $T+1$ and all other shocks being zero. Our counterfactual is given by the forecast $\widetilde{\boldsymbol{y}}^{cf}_{T+1,T+h}$ conditional on a global uncertainty shock materialising in period $T+1$ and the additional constraint that the dollar exchange rate does not change over the forecast horizon. The latter is achieved through the materialisation of US monetary policy shocks that offset the effects of the global uncertainty shock on the dollar exchange rate.[21]

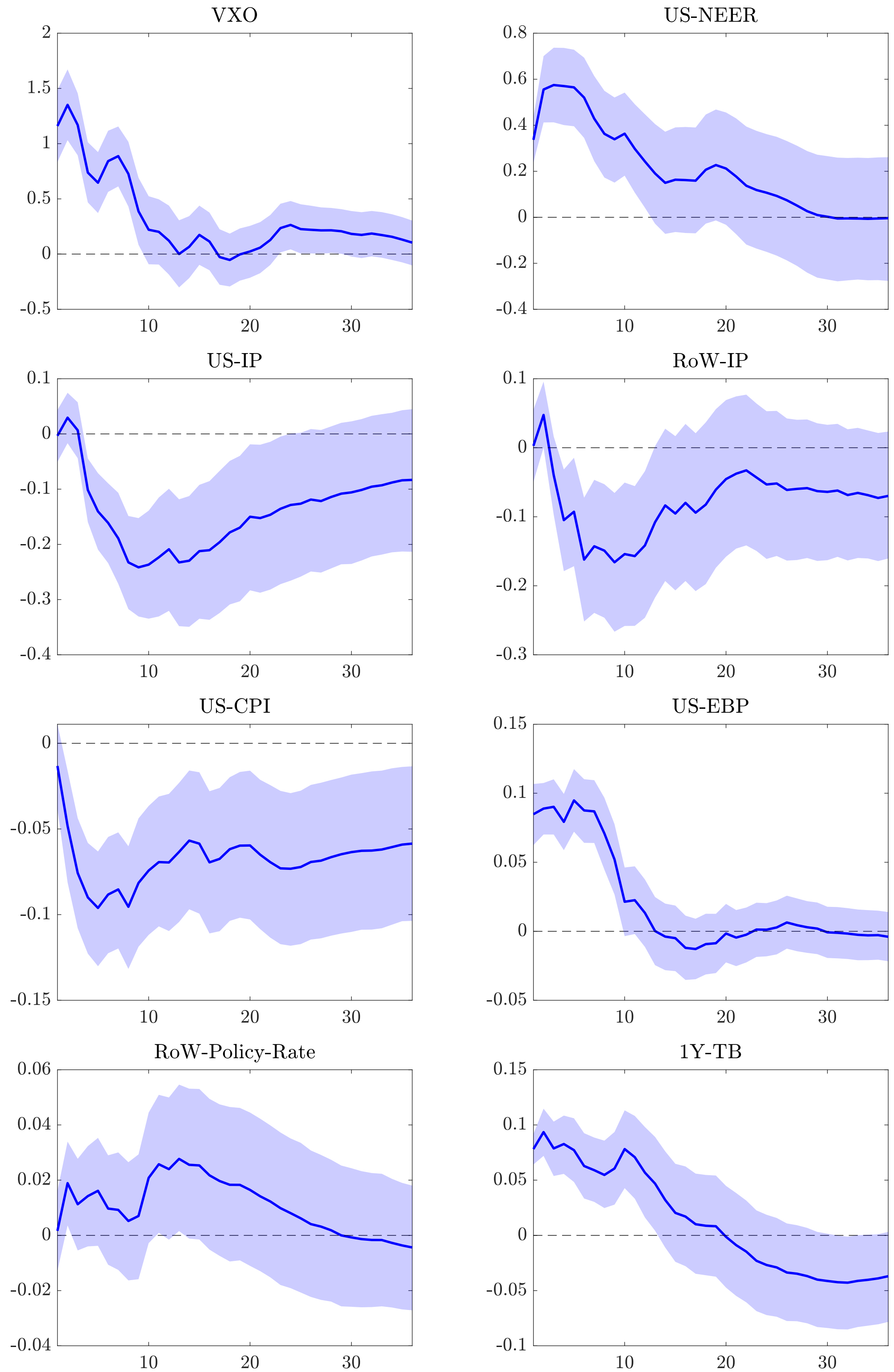
4.3.2 Results from structural shock no-appreciation counterfactuals

Before showing the results of the SSC, we first present the effects of a US monetary policy shock. Figure 10 shows the impulse responses to a US monetary policy shock identified using the intra-daily interest rate surprises net of central bank information effects as proxy variable in the BPSVAR mdoel; recall that we identify the US monetary policy shock and the global uncertainty shock jointly.

The results suggest that a contractionary US monetary policy shock induces a tightening in US financial conditions as reflected in the rise of the excess bond premium as in Gertler & Karadi (2015) and the VXO as in (Bekaert et al. 2013; Rey 2016; Miranda-Agrippino & Rey 2020). The dollar appreciates on impact, and again exhibits some delayed overshooting, even if it is rather short

[21] See Breitenlechner et al. (2020) for further technical details and the specification of the matrices $\overline{\boldsymbol{C}}$, $\overline{\boldsymbol{f}}_{T+1,T+h}$, $\boldsymbol{\Xi}$, $\overline{\boldsymbol{g}}_{T+1,T+h}$, $\boldsymbol{\Omega}_g$ and $\overline{\boldsymbol{\Omega}}_f$ under the baseline and the counterfactual conditional forecast.

Figure 10: Responses to a contractionary US monetary policy shock



Note: The figure presents the impulse responses to a one-standard deviation US monetary policy shock. See the notes to Figure 2.

relative to that documented in Eichenbaum & Evans (1995).[22] US consumer prices fall persistently. The contractionary US monetary policy shock causes a globally synchronized contraction, consistent with existing literature (Georgiadis 2016; Dedola et al. 2017; Dees & Galesi 2019; Iacoviello & Navarro 2019; Degasperi et al. 2020). Interestingly, despite the contraction, monetary policy in RoW also tightens.[23]

Figure 11 presents the results from the SSC under which US monetary policy responds—in contrast to the regularities in the data over our sample period—to the global uncertainty shock so as to prevent dollar appreciation. As in Figure 4, the solid blue lines represent response of the global uncertainty shock from the baseline, and the red circled lines represent the SSC counterfactual. In the counterfactual, US monetary policy is loosened much more strongly than in the baseline. In fact, the reduction in the US policy rate in the counterfactual is more than twice as large as in the baseline. This additional US monetary policy loosening prevents the appreciation of the dollar exchange rate. Interestingly, the additional loosening also prevents an increase in the VXO. Hence, by stabilising the dollar exchange rate, US monetary policy effectively undoes the effects of the global uncertainty shock on financial market volatility as measured by the VXO. US consumer prices also remain roughly constant in the counterfactual, while the excess bond premium even slightly decreases. Despite the additional US monetary policy loosening, there is still a fall in world industrial production. However, the slowdown in real activity in the counterfactual is muted relative to the baseline, both in the US and RoW. In sum, we find that if US monetary policy loosened more aggressively than it typically has been in the data it could mitigate substantially the contractionary effects of global uncertainty shocks, both in the US and RoW.

5 Conclusion

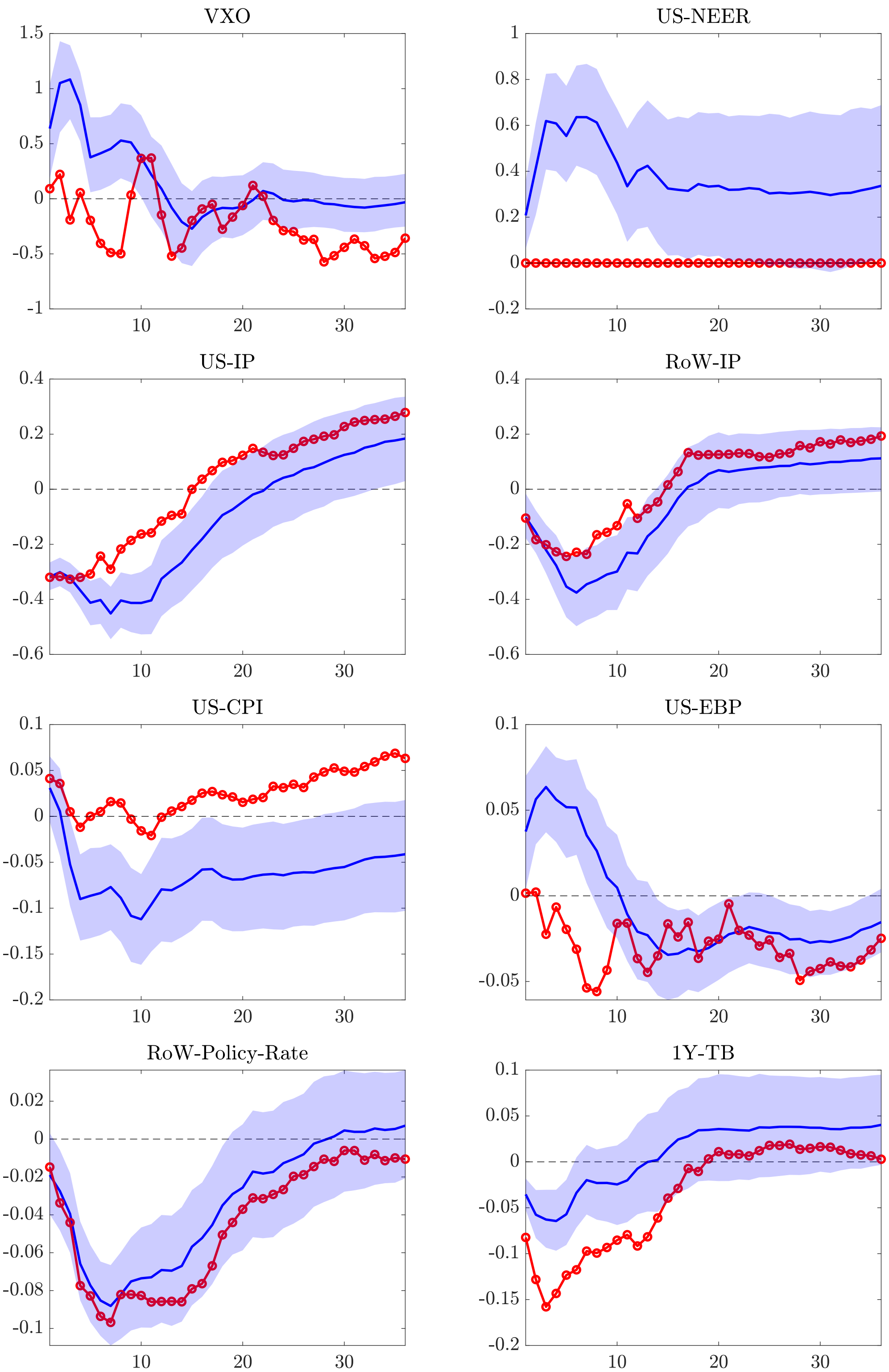
In this paper we document that global uncertainty shocks cause a slowdown in global real activity, a tightening in global financial conditions, and, in particular, an appreciation of the US dollar exchange rate. Conventional wisdom suggests that the dollar appreciation mitigates the negative output effects outside the US due to expenditure switching. However, recent research suggests dollar appreciation might amplify the effects of a global uncertainty shock on financial conditions and real activity through a financial channel of exchange rates that operates on cross-border US dollar liabilities. The contribution of the dollar appreciation to the overall effects of global uncertainty shocks is thus ambiguous in theory.

We analyse empirically which of the two effects dominates. In particular, we explore counterfactual scenarios in which the dollar exchange rate does not respond to the global uncertainty shock. We find that the financial channel of exchange rates dominates expenditure switching: In the absence of dollar appreciation, the tightening in global financial conditions and the slowdown in real activity

[22] The adjustment of the dollar exchange rate in Figure 10 is in line with the notion that the delay may be shorter in more recent samples (Kim et al. 2017) or be due to differences in the identification strategy.

[23] That RoW mirror US policy rates is consistent with fear-of-floating, both for AEs and EMEs (Calvo & Reinhart 2002; Corsetti et al. 2021; Georgiadis & Zhu 2021). Degasperi et al. (2020) also find that RoW policy rates rise in response to a contractionary US monetary policy shock.

Figure 11: Baseline and SSC-based counterfactual responses to a global uncertainty shock



Note: The circled red lines depict the SSC-based counterfactual responses to a global uncertainty shock.

is less pronounced than in the baseline; while the response of US net exports barely changes if dollar appreciation is absent, cross-border US dollar denominated credit drops by much less than. Finally, we show that dollar appreciation could be prevented if US monetary policy responded more accommodatively, eventually dampening the effects of global uncertainty shocks. An important issue that we do not address in this paper is the possibility of an asymmetric relationship between uncertainty shocks one the one hand and the VXO and the dollar exchange rate on the other hand. We leave this for future research.

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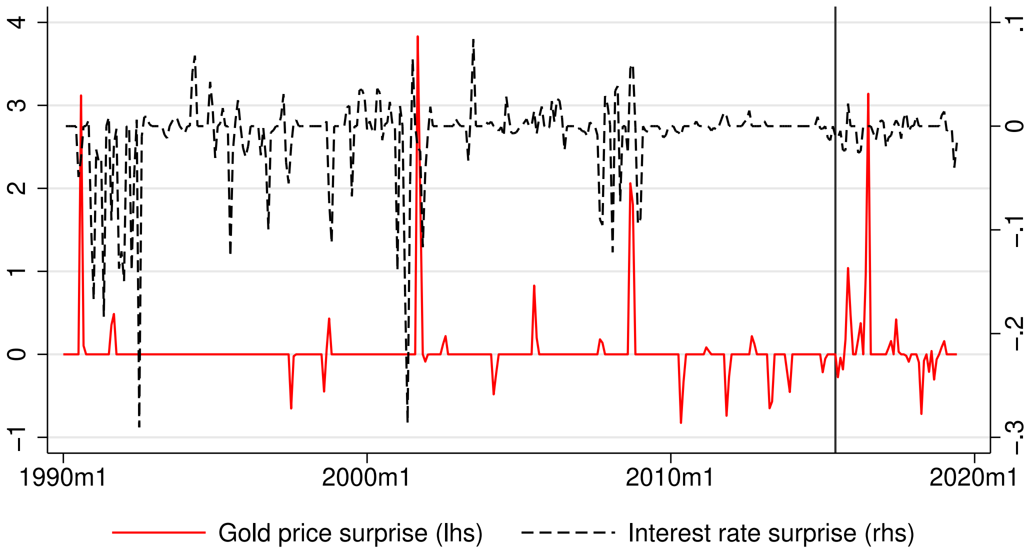
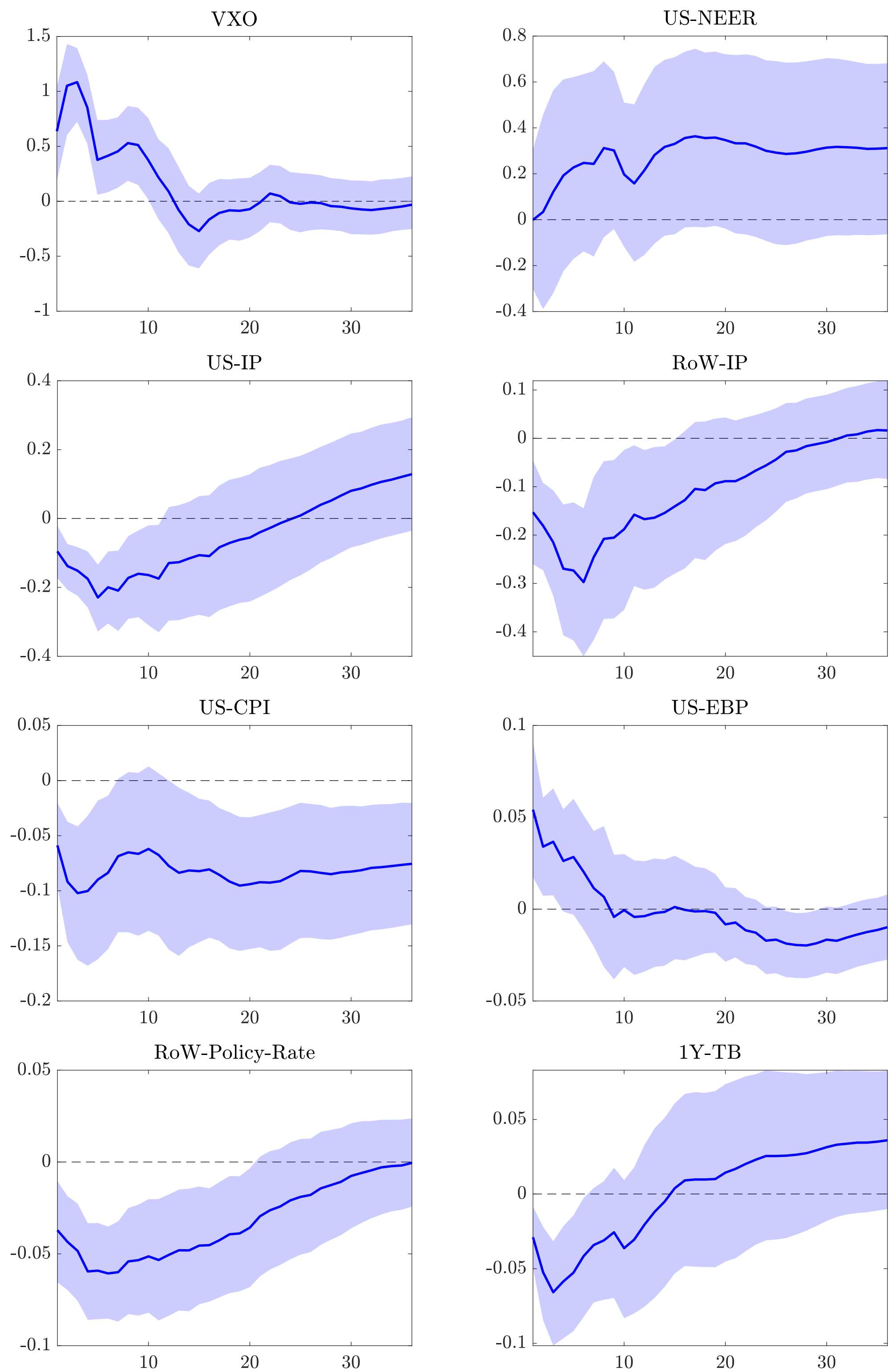
A Additional figures

Figure A.1: Monthly time series of gold price and interest rate surprises



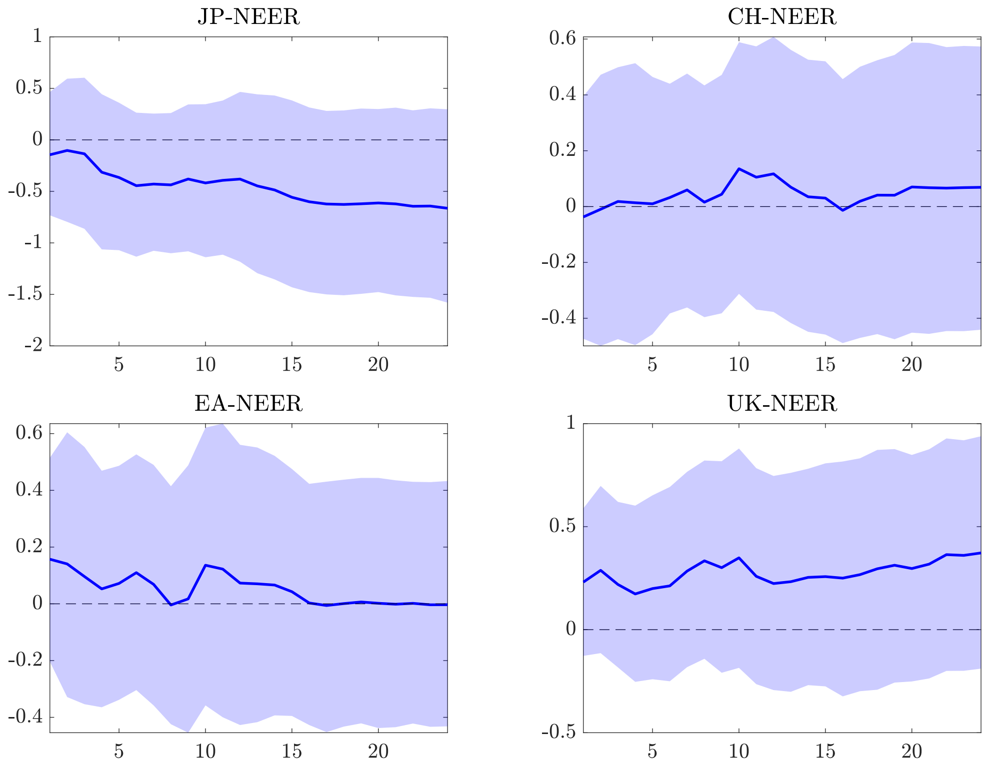
Note: The figure shows the time series of the gold price (red solid line) and 3-month Federal Funds Futures rate (black ashed line) surprises of Piffer & Podstawski (2018) as well as Jarociński & Karadi (2020) in percent and at the monthly frequency. The vertical line in 2015 indicates the data after which the narratively selected dates for the global uncertainty shock events are taken from Bobasu et al. (2020). We aggregate the daily changes in the proxy variables to monthly frequency as in Gertler & Karadi (2015), that is we first create a cumulative daily surprise series, then, second, take monthly averages of these series, and, third, obtain monthly average surprises as the first difference of this series.

Figure A.2: Impulse responses to global demand shock



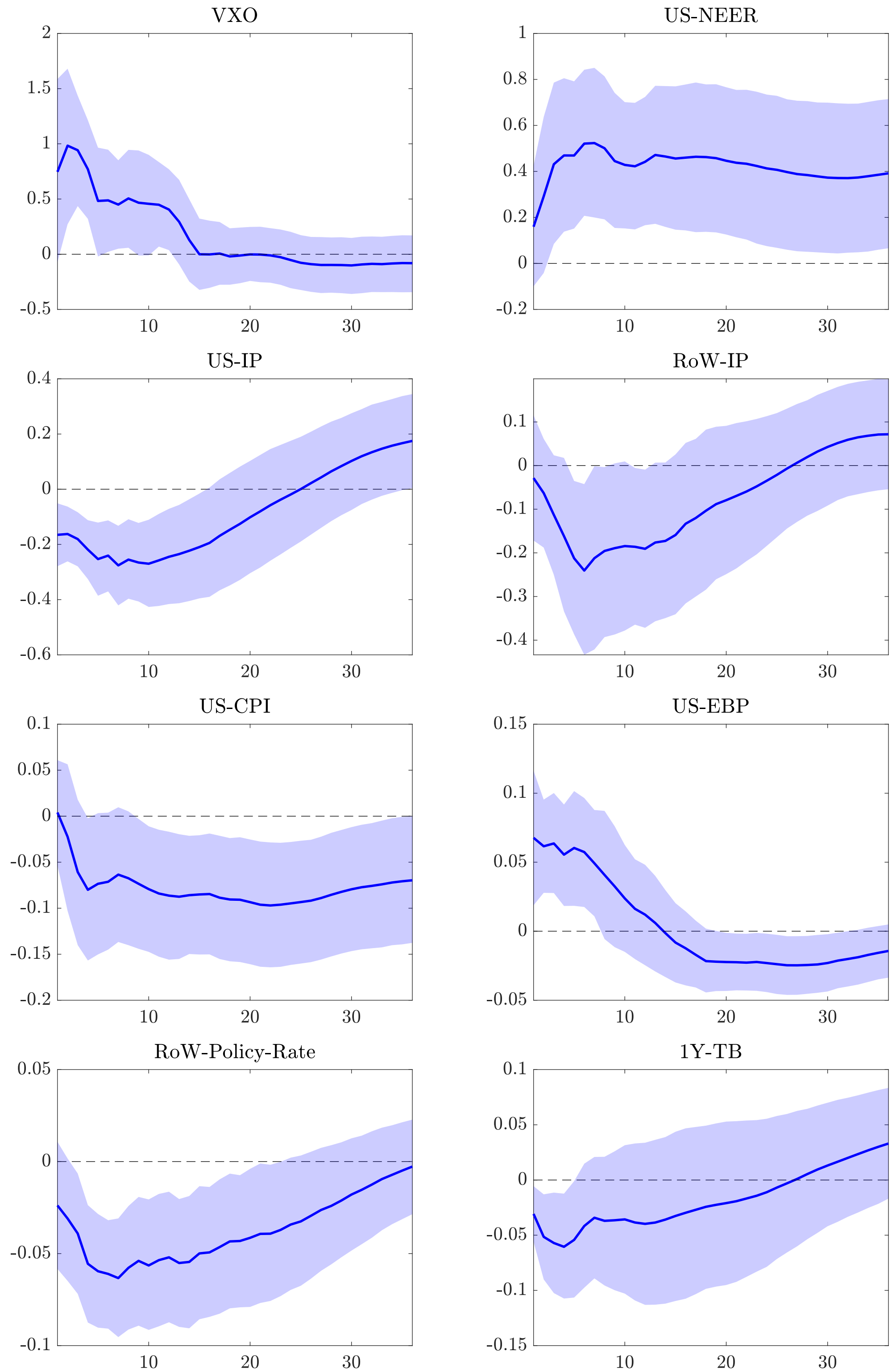
Note: The figure presents the impulse responses to a one-standard deviation global demand shock identified based on sign restrictions. See also the notes to Figure 2.

Figure A.3: Responses of the Japanese yen, Swiss franc, euro and pound sterling exchange rates to a global demand shock



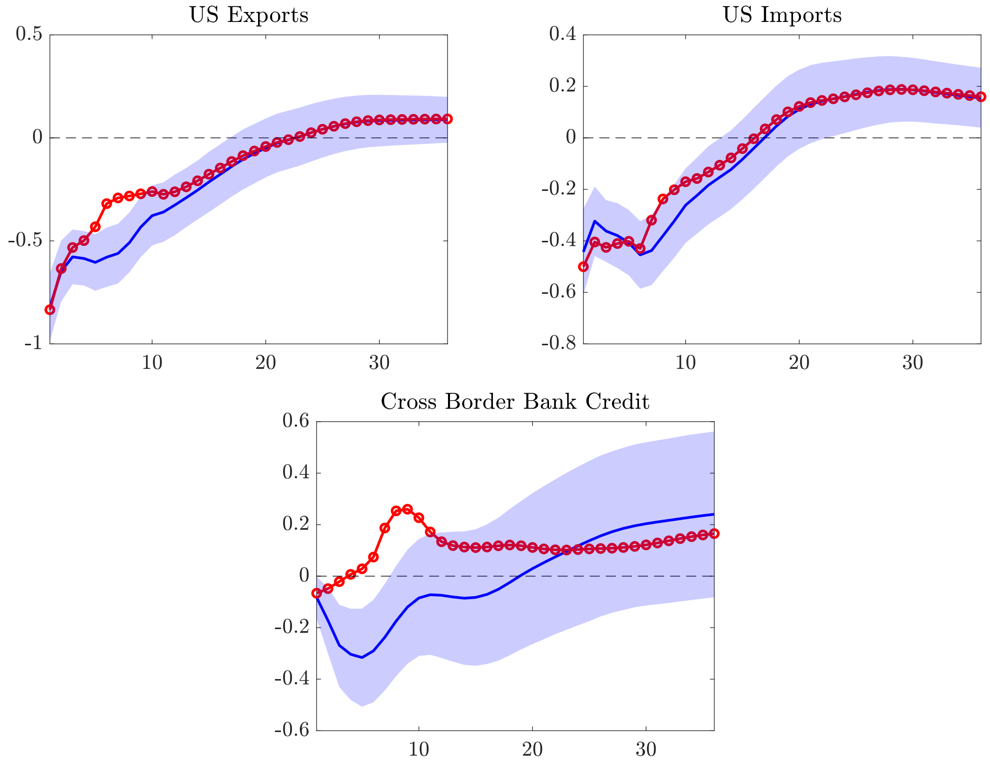
Note: The figure presents the impulse responses to a one-standard deviation global demand shock identified based on sign restrictions. See also the notes to Figure 2.

Figure A.4: Impulse responses to global uncertainty shock when allowing the gold price surprises to be correlated with all structural shocks



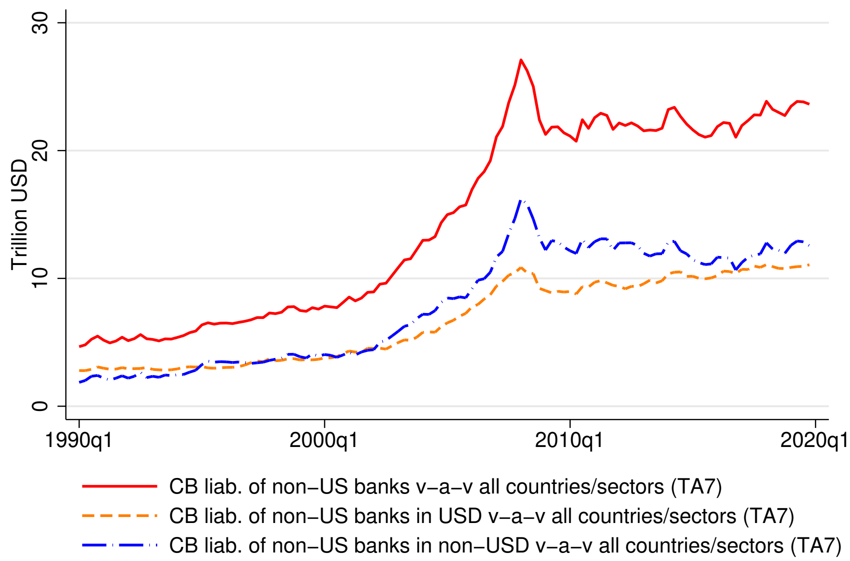
Note: The figure presents the impulse responses to a one-standard deviation global uncertainty shock based on an alternative identification scheme in which the gold price surprises are allowed to be correlated with all structural shocks, imposing only that the correlation is strongest with the global uncertainty shock. See also the notes to Figure 2.

Figure A.5: Baseline and MRE-based counterfactual responses of cross-border bank credit and US trade to a global uncertainty shock from a large BPSVAR



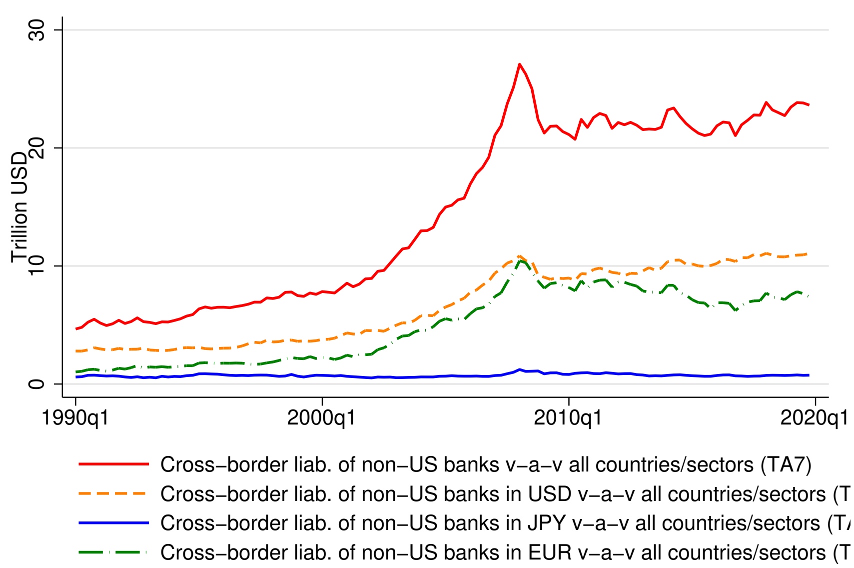
Note: See the notes to Figure 4. The results are based on a BPSVAR model that includes US exports and imports as well as cross-border credit simultaneously. The model is estimated with informative Minnesota-type priors and optimal hyperpriors/prior tightness as in Giannone et al. (2015).

Figure A.6: Evolution of cross-border bank credit: Summary



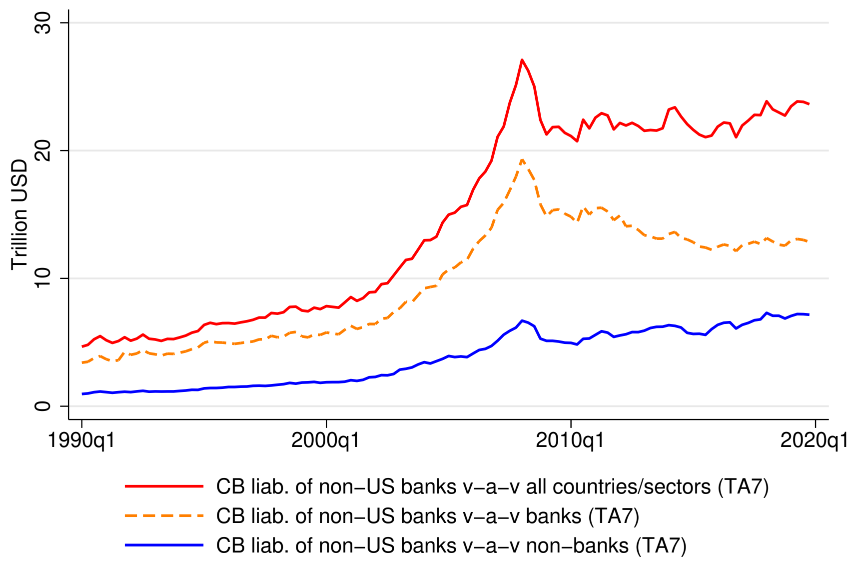
Note: The figures presents the evolution of cross-border bank credit from the BIS Locational Banking Statistics. See Table A.1 for variable definitions.

Figure A.7: Evolution of cross-border bank credit by currencies



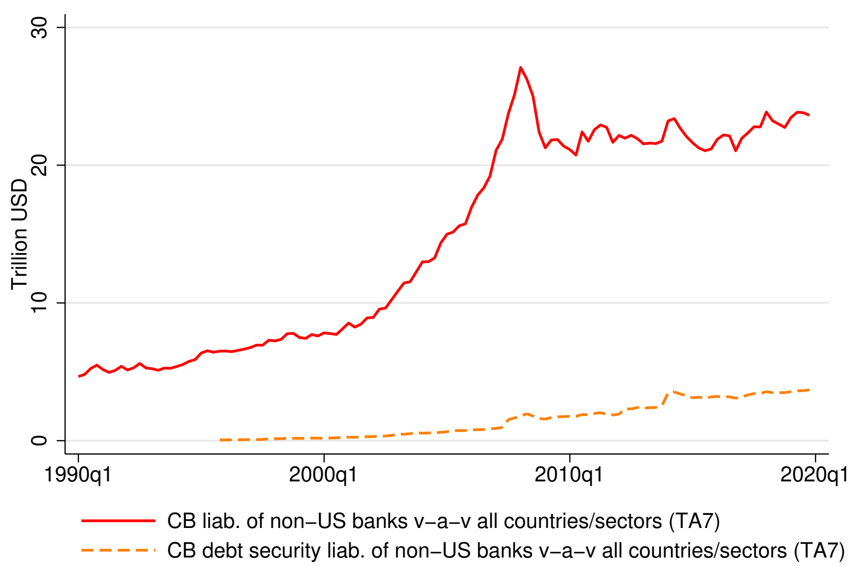
Note: The figures presents the evolution of cross-border bank credit from the BIS Locational Banking Statistics. See Table A.1 for variable definitions.

Figure A.8: Evolution of cross-border bank credit by creditor type



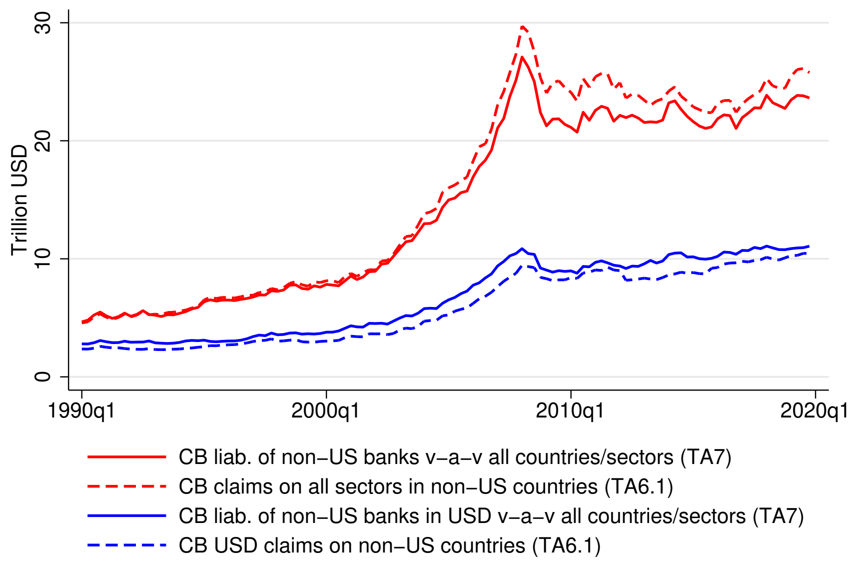
Note: The figures presents the evolution of cross-border bank credit from the BIS Locational Banking Statistics. See Table A.1 for variable definitions.

Figure A.9: Evolution of cross-border bank credit by instrument



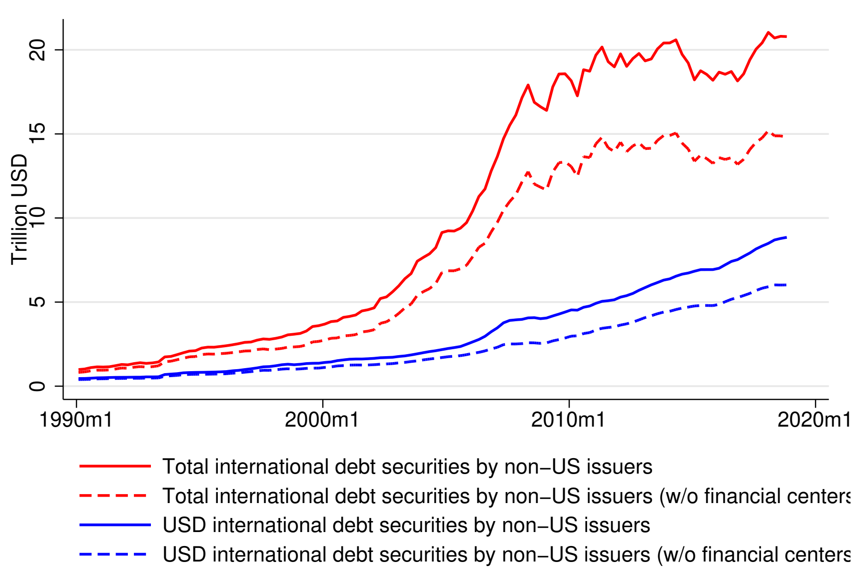
Note: The figures presents the evolution of cross-border bank credit from the BIS Locational Banking Statistics. See Table A.1 for variable definitions.

Figure A.10: Comparison of cross-border bank credit claims and liabilities



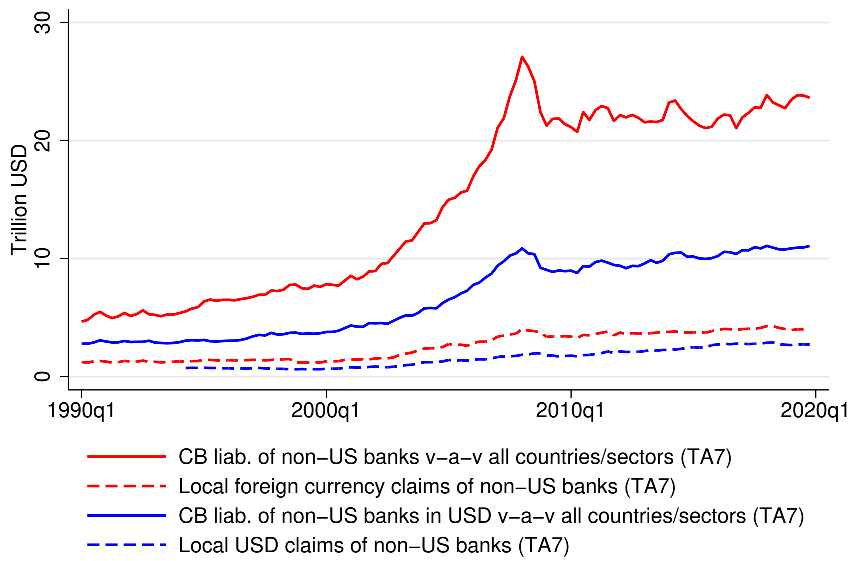
Note: The figures presents a comparison of the cross-border bank credit data from Tables A7 and A6.1 from the BIS Locational Banking Statistics. Information is only available for BIS reporting banks, so cross-border claims and liabilities of reporting banks need not coincide. See Table A.1 for variable definitions.

Figure A.11: Evolution of international debt securities



Note: The figures presents the evolution of international debt securities. The list of financial centers is taken from Bertaut et al. (2019). See Table A.1 for variable definitions.

Figure A.12: Evolution of local credit of globally active banks



Note: The figures presents the evolution of local claims of globally active banks. See Table A.1 for variable definitions.

Table A.1: Data description

Variable	Description	Source	Coverage
US 1-year TB rate	1-year Treasury Bill yield at constant maturity	US Treasury/Haver	1990m1 - 2019m6
US IP	Industrial production excl. construction	FRB/Haver	1990m1 - 2019m6
US CPI	US consumer price index	BLS/Haver	1990m1 - 2019m6
US EBP		Favara et al. (2016)	
US dollar NEER	Nominal broad trade-weighted Dollar index	FRB/Haver	1990m1-2019m6
VXO	CBOE market volatility index VXO	Wall Street Journal/Haver	1990m1 - 2019m6
RoW IP	Industrial production, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m6
RoW policy rate	Short-term official/policy rate, see Martínez-García et al. (2015)	Dallas Fed Global Economic Indicators/Haver	1990m1 - 2019m6
Yen, euro, Swiss franc, British pound NEER	Nominal broad effective exchange rate	J.P. Morgan/Haver	1990m1-2019m6
US exports	Exports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
US imports	Imports of goods and services (chnd. 2012$)	BEA/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US USD cross-border bank credit	Banks' external liabilities in USD of banks owned by the world less external liabilities in USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
Non-US non-USD cross-border bank credit	Banks' external liabilities in non-USD of banks owned by the world less external liabilities in non-USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
USD local banking claims	Local claims in USD of banks owned by the world less local claims in USD of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
EUR/Yen local banking claims	Local claims in EUR/Yen of banks owned by the world less local claims in EUR/Yen of banks owned by US nationals	BIS Locational Banking Statistics, Table A7/Haver	1990q1-2019q2, interpolated to monthly frequency
USD international debt securities	USD debt securities issued by non-residents	BIS International Debt Issuance Statistics/Benetrix et al. (2020)	1990q1-2018q4, interpolated to monthly frequency
Non-USD international debt securities	non-USD debt securities issued by non-residents	BIS International Debt Issuance Statistics/Benetrix et al. (2020)	1990q1-2018q4, interpolated to monthly frequency

Notes: BLS stands for Bureau of Labour Statistics, FRB for Federal Reserve Board, BEA for Bureau of Economic Analysis, and BIS for Bank for International Settlements.